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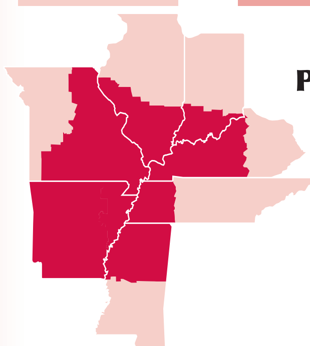

Proceedings of the First Annual Conference of the Business & Economics Research Group (BERG)

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Contributing Authors

Nan-Ting Chou
University of Louisville
ntchou01@gwise.louisville.edu

Paul Coomes
University of Louisville
paul.coomes@louisville.edu

Bruce Domazlicky
Southeast Missouri State University
bdomazlicky@semovm.semo.edu

John E. Garen
University of Kentucky
jgaren@uky.edu

William H. Hoyt
University of Kentucky
whoyt@uky.edu

Michael T. Owyang
Federal Reserve Bank of St. Louis
Michael.T.Owyang@stls.frb.org

David A. Penn
Middle Tennessee State University
dpenn@mtsu.edu

Jeremy M. Piger
Federal Reserve Bank of St. Louis
Jeremy.M.Piger@stls.frb.org

David E. Rapach
Saint Louis University
rapachde@slu.edu

John P. Shelnutt
Arkansas Department of Finance and Administration
John.Shelnutt@dfa.state.ar.us

Anna L. Stewart
University of Kentucky
alstew01@email.uky.edu

Jack K. Strauss
Saint Louis University
strausjk@slu.edu

Howard J. Wall
Federal Reserve Bank of St. Louis
Howard.J.Wall@stls.frb.org

Vincent W. Yao
University of Arkansas at Little Rock
wxyao@ualr.edu

Editor's Introduction

Howard J. Wall

BRANCHING OUT

The Federal Reserve Bank of St. Louis has recently embarked on an initiative called Branching Out, which will increase the Bank's participation in the geographic region of the Federal Reserve System's Eighth District. Specifically, the St. Louis Fed will make greater contributions to the development of the economies of its branch cities—Little Rock, Louisville, and Memphis—and the surrounding region.

CRE8

As part of this initiative, the Research Division of the St. Louis Fed established its Center for Regional Economics–8th District (CRE8) in January 2005. CRE8 will provide and facilitate rigorous economic analysis of policy issues affecting local, state, and regional economies—particularly those in the Eighth District. In addition to producing its own research, CRE8 will organize policy forums, conferences, and symposia that highlight economic research done outside the St. Louis Fed. These events will inform and initiate discussion among policymakers in our region's communities.

BERG

The goals of Branching Out are also being accomplished through another channel: the newly established consortium known as the Eighth District Business and Economics Research Group (BERG). BERG is composed of CRE8 and university-based centers for business and economics research in the states of the Eighth District:

Institute for Economic Advancement, University of Arkansas at Little Rock

Center for Business and Economic Research, University of Arkansas at Fayetteville

Center for Business Development and Economic Research, Jackson State University, Mississippi

Sparks Bureau of Business and Economic Research, University of Memphis

Center for Business and Economic Research, University of Tennessee, Knoxville

Center for Business and Economic Research, University of Kentucky

Economic and Policy Analysis and Research Center, University of Missouri at Columbia

Louisville Economic Monitor, University of Louisville

Center for Economic and Business Research, Southeast Missouri State University

Delta Center for Economic Development, Arkansas State University

Business and Economic Research Center, Middle Tennessee State University

Simon Center for Regional Forecasting, Saint Louis University

BERG's purpose is to create a forum for researchers who have a detailed knowledge of the sub-areas of the Eighth District. The St. Louis

Howard Wall is an assistant vice president and the director of the Center for Regional Economics–8th District (CRE8) at the Federal Reserve Bank of St. Louis.

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Fed will benefit from BERG's valuable sources of information on the economic trends and conditions within the District. BERG members will benefit from the network of organizations and its presentations, discussions, and comparisons of state- and local-level research.

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The primary channel for organizing and distributing the research generated from these varied sources and events is our newly formed journal, *Regional Economic Development*. With the CRE8 staff as the editorial board, this journal will publish the proceedings from these events on the St. Louis Fed's Research Division web site: http:research.stlouisfed.org/publications. Although the contents of the journal will be accessible to a wide audience, beyond purely academic circles, it will nevertheless offer serious economic analysis, addressing and solving practical policy issues.

Upcoming events whose proceedings will be published in *Regional Economic Development* include a symposium on November 4, 2005, which CRE8 is co-hosting with the Wiedenbaum Center at Washington University in St. Louis, titled "Challenges to Public Education Financing Facing Missouri and the Nation." Also, on March 29 and 30, 2006, CRE8 will co-host "TED2006," a conference on transportation and economic development organized by the Institute for Economic Advancement at the University of Arkansas at Little Rock.

This inaugural issue of *Regional Economic Development* contains the proceedings of the first annual conference of BERG, held in St. Louis on May 6, 2005. The first three articles of this issue discuss recent and ongoing trends in the Eighth District. Jeremy Piger, Michael Owyang, and Howard Wall of the St. Louis Fed look at the recent business cycle experience of the seven states that lie wholly or partly in the District (Arkansas, Illinois, Indiana, Kentucky, Mississippi, Missouri, and Tennessee). Paul Coomes and Nan-Ting Cho of the University of Louisville use several different indicators to describe the experience of the Louisville economy since 1990, while David Penn of Middle Tennessee State University discusses the booming economy of Middle Tennessee and the challenges faced by local governments in expanding their services to keep up with growth.

The fourth and fifth articles in this volume are both concerned with income inequality. Bruce Domazlicky of Southeast Missouri State University describes how income inequality decreased in southeast Missouri between 1990 and 2000 and estimates the effects of various factors in determining the level of inequality. Cross-county variation in income inequality in Arkansas is the focus of the article by John Shelnutt of the Arkansas Department of Finance and Administration and Vincent Yao of the Institute for Economic Advancement of the University of Arkansas at Little Rock. They also compare the different sets of variables that explain differences in inequality in MSAs versus non-MSAs.

The penultimate article, by William Hoyt, John Garen, and Anna Stewart of the University of Kentucky, compares the cost of government services in Kentucky to that of neighboring states. Among their many findings is that there are substantial economies of scale in the provision of government services and that the importance of scale economies differs across types of government service. In the last article, David Rapach and Jack Strauss of Saint Louis University examine a variety of different approaches to forecasting employment growth in Missouri. They conclude that simple combining methods offer a low-cost way of generating reliable forecasts.

I would like to acknowledge the help and suggestions of the authors, referees, and conference participants in putting together this conference volume. In addition, I would like to thank the managing editor, George Fortier, and his staff, Lydia Johnson and Donna Stiller, for their fine work in putting the issue together in record time. Finally, I would like to thank Sandra Butler for all of her help in organizing the conference.

The 2001 Recession and the States of the Eighth Federal Reserve District

Michael T. Owyang, Jeremy M. Piger, and Howard J. Wall

This paper examines and compares the recent business cycle experiences of the seven states that lie partly or wholly within the Eighth Federal Reserve District (Arkansas, Illinois, Indiana, Kentucky, Mississippi, Missouri, and Tennessee). For the period surrounding the 1990-91 recession, six of the seven states had recessions that were much shorter than that for the country as a whole. In addition, for the period surrounding the 2001 recession, four states (Arkansas, Indiana, Kentucky, and Tennessee) entered and exited recession earlier than the country as a whole. Recessions in the other three states began earlier and ended later than the recession for the country as a whole.

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This paper examines and compares the recent business cycle experiences of the seven states that lie partly or wholly within the Eighth Federal Reserve District (Arkansas, Illinois, Indiana, Kentucky, Mississippi, Missouri, and Tennessee). We pay particular attention to the period surrounding the 2001 national recession. Our analysis relies on the supposition that state-level business cycles can be characterized as a series of distinct recession and expansion phases, as is commonly held to be true of the national business cycle. The primary example of such a characterization of the national business cycle is the activity of the National Bureau of Economic Research (NBER) Business Cycle Dating Committee, which provides semiofficial recession and expansion dates.

Because the NBER chronology is available only for U.S. national economic activity, alternative methods must be used to identify business cycle turning points in state-level data. To this end, we follow a recent paper by Owyang, Piger, and Wall (2005), hereafter simply OPW, in estimating state-level turning points with a version of the regime-switching model of Hamilton (1989). As with the NBER, the Hamilton model is based on the notion that the business cycle can be split into distinct recession and expansion phases. In fact, the Hamilton model can be thought of as a nonjudgmental, statistical alternative to the committee-consensus method of the NBER.

A significant hurdle in determining business cycle turning points at the state level is the inadequacy of data relative to what is available for the national economy. When applied to the national economy, the Hamilton model is typically applied to gross domestic product (GDP), which has a quarterly frequency and has been found to provide distinct turning points in and out of expansion and recession phases.[1] At the state level, however, the analog to GDP—gross state product—is produced only with an annual frequency and with a 2- to 3-year lag, making it of little use in detecting phase shifts. The solution in OPW is to use the state-level coincident index (SCI) produced by the Federal Reserve Bank of Philadelphia and described by Crone and Clayton-Matthews (2005).

[1] Boldin (1994) and Chauvet and Piger (2003), among others, have found that the Hamilton model does quite well in mimicking the turning point dates of the NBER.

Michael T. Owyang and Jeremy M. Piger are senior economists at the Federal Reserve Bank of St. Louis. Howard Wall is an assistant vice president and the director of the Center for Regional Economics–8th District (CRE8) at the Federal Reserve Bank of St. Louis.

The SCI uses the dynamic factor model of Stock and Watson (1989) to combine four variables—the unemployment rate, payroll employment, average weekly manufacturing hours, and real wage and salary disbursements—into a single indicator of state-level labor-market activity.

The main advantage of the SCI is that, as demonstrated by OPW, it provides timely and frequent observations of series that tend to have distinct state-level business cycle turning points. Further, it provides a much cleaner and better-behaved variable than any of its components, which are much noisier and more erratic than their national-level counterparts. The disadvantage of the SCI is that, because it uses labor-market variables only, it is not as broad a measure of activity as GDP. As such, it is probably best viewed as an indicator of overall state-level labor-market conditions rather than as a coincident indicator of gross state product or some other broad measure of state-level conditions. With this in mind, we also apply the Hamilton model to national nonfarm payroll employment to provide more-relevant national recession and expansion dates to compare with our state-level dates.

MODEL AND ESTIMATION

In the Hamilton (1989) Markov-switching model, different business cycle phases are treated as arising from different models, each with its own mean growth rate. Let μ_0 be the mean growth rate when the economy is in expansion, and let μ_1, which is normalized to be negative, be the difference between the mean growth rates in expansion and recession. Specify a simple model for the growth rate of some measure of economic activity, y_t, as

$$y_t = \mu_0 + \mu_1 S_t + \varepsilon_t, \qquad \mu_1 < 0. \tag{1}$$

To introduce recession and expansion phases, the mean growth rate in (1) switches between the two regimes, where the switching is governed by a state variable, $S_t = \{0,1\}$. Deviations from this mean growth rate are created by the stochastic disturbance, $\varepsilon_t \sim N(0,\sigma_\varepsilon^2)$. When S_t switches from 0 to 1, the growth rate switches from μ_0 to $\mu_0 + \mu_1$. Because $\mu_1 < 0$, S_t switches from 0 to 1 at times when the economy switches from the high-growth phase to the low-growth phase (expansion to recession) or vice versa.

The switching variable, S_t, is unobserved, meaning that we need to place restrictions on the probability process governing it. We assume that the process for S_t is a first-order two-state Markov chain, implying that any persistence in the regime is completely summarized by the value of the state in the previous period. More specifically, the probability process driving S_t is captured by the transition probabilities $\Pr[S_t = j \mid S_{t-1} = i] = p_{ij}$. We estimate the model using the multi-move Gibbs-sampling procedure for Bayesian estimation of Markov-switching models implemented by Kim and Nelson (1999).[2]

Our data are monthly observations of the SCIs over the period 1983-2004 for the seven states of the Eighth District. We restrict our estimation to post-1982 data to avoid possible problems with structural breaks. Clarida, Galí, and Gertler (2000) and Boivin and Giannoni (2003), for example, show how monetary policy shocks have much smaller effects on output in the post-Volcker period. Also, McConnell and Perez-Quiros (2000) demonstrate how national output growth has been significantly less volatile since the early 1980s.

Our first step is to use the Hamilton model to obtain a useful description of the national business cycle that we can compare with the state-level business cycles from the SCIs. As we have mentioned, because the SCIs are indicators of labor-market conditions, we use national nonfarm payroll employment to describe the national employment cycle, which grew at an average monthly rate of 0.15 percent during our sample period.[3] For reference, monthly growth of payroll

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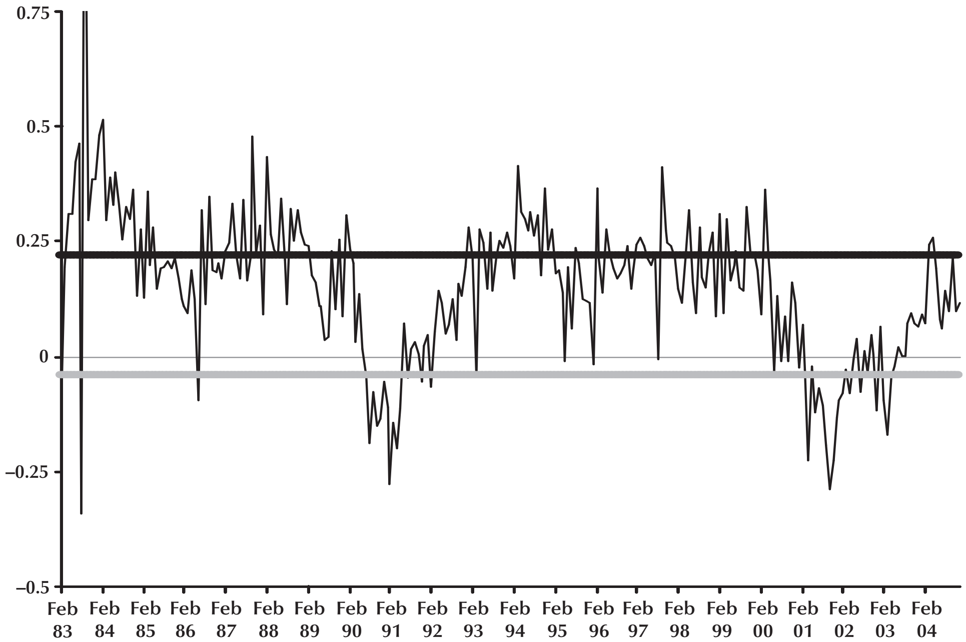
[2] The Gibbs sampler draws iteratively from the conditional posterior distribution of each parameter, given the data and the draws of the other parameters of the model. These draws form an ergodic Markov chain whose distribution converges to the joint posterior distribution of the parameters, given the data. When we simulate the posterior distribution, we discard the first 2,000 draws to ensure convergence. Descriptive statistics regarding the sample posterior distributions are then based on an additional 10,000 draws.

[3] We should note that the employment series from the household survey is of limited use for our purposes because its monthly growth does not exhibit the distinct breaks found in payroll employment.

Figure 1

U.S. Payroll Employment Growth (percent)



NOTE: The thick black line is the average expansion growth rate; the thick gray line is the average recession growth rate.

employment over the period was 0.18 in Arkansas, 0.10 in Illinois, 0.15 in Indiana, 0.18 in Kentucky, 0.14 in Mississippi, 0.13 in Missouri, and 0.18 in Tennessee.

According to the model, this average growth rate is the average of the recession and expansion growth rates weighted by the frequencies of the two business cycle phases. The Hamilton model provides estimates of the average growth rates in each of the two phases and, for each observation, the probability that the labor market is in the recession phase. Applying the Hamilton model yields average monthly national employment growth rates of 0.22 percent during expansion and –0.04 percent during recession.[4]

In divining the probability of recession, the model compares the actual growth rate to the average growth rates for the two phases. The model also considers how persistently this relative proximity is. Figure 1 shows actual average U.S. employment growth relative to the average growth rates for the two phases. The probability that the national labor market is in recession is provided by Figure 2, where the shaded areas indicate periods of national NBER recessions.[5] As Figure 2 shows, the model does a good job of separating the data into recession and expansion phases in that the probability of recession rises and falls rapidly as the national labor market switches between phases. The only period for which the model provides mixed signals is for August 1983 (see Figure 1). That month appears to be anomalous in that employment growth spiked down in August, only to spike up in September to more than make up for the previous month's job losses.

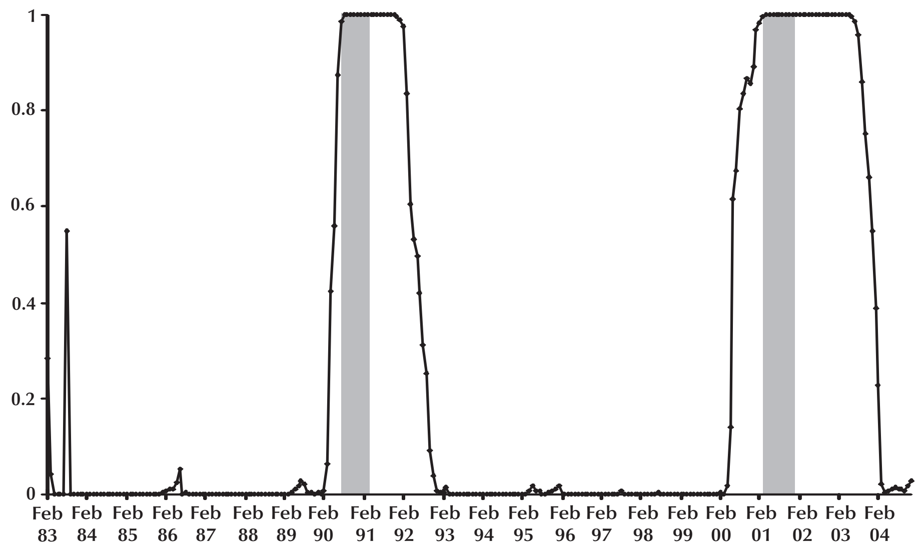
The main result apparent in Figure 2 is the different timings and lengths of the NBER and

[4] The 90 percent coverage intervals for these growth rates are (0.20, 0.23) and (–0.07, –0.01), respectively.

[5] According to the NBER, the national economy was in recession twice during the sample period: August 1990–March 1991 and April 2001–November 2001.

Figure 2

Probability of U.S. Employment Recession



NOTE: Shaded areas indicate NBER recessions: August 1990–March 1991 and April 2001–November 2001. National labor-market recessions: June 1990–April 1992 and June 2000–November 2003.

national labor-market recessions. Using an admittedly arbitrary recession probability of 0.6 or higher to indicate months of recession, the first national labor-market recession began in June 1990 (2 months before the start of the NBER recession) and ended in April 1992 (more than a year after the end of the NBER recession). This long lag between the ends of the NBER and labor-market recessions was the period of the so-called jobless recovery.

The 2001 NBER recession shared some of the features of the 1990-91 NBER recession. In particular, it did not last nearly as long as the associated labor-market recession did, which began much earlier and ended much later than the 2001 NBER recession: The national labor market entered recession in June 2000—10 months prior to the April 2001 start of the NBER recession. Further, although the NBER recession ended in November 2001, it was not until a full 2 years later that the national labor market saw an end to its recession. Thus, the disjointedness between the labor-market and the broader economy was significantly greater with the 2001 NBER recession than the 1990-91 NBER recession: In 2001, the national economy went into recession well after the national labor market did,[6] and the jobless recovery lasted almost twice as long as it did in 1991-92.

ESTIMATION OUTPUT FOR THE STATES

Growth Rates

The estimated state-level average monthly growth rates in expansion and recession are provided in Table 1, along with the actual growth rates for the period 1983-2004. Of the seven states of the Eighth District, Tennessee had the highest average growth rate over the sample period

[6] We should note that this overstates the difference between the start of the NBER recession and the surrounding national labor-market recession. The NBER determined the start of the recession before data for 2000 were revised downward significantly. Using revised versions of the same data used by the NBER, Chauvet and Piger (2005) determine that the national economy entered recession in November 2000.

Table 1

Actual and Estimated Growth Rates, 1983-2004 (percent)

	Average monthly growth rate 1983-2004	Monthly growth rate in expansion	Monthly growth rate in recession	Difference between expansion and recession growth rates
Arkansas	0.29	0.36 (0.34, 0.41)	0.02 (–0.04, 0.13)	0.34
Illinois	0.27	0.41 (0.39, 0.43)	–0.07 (–0.12, –0.03)	0.49
Indiana	0.30	0.39 (0.37, 0.42)	–0.11 (–0.17, –0.06)	0.51
Kentucky	0.29	0.36 (0.35, 0.38)	–0.03 (–0.07, 0.02)	0.39
Mississippi	0.25	0.34 (0.32, 0.36)	0.02 (–0.03, 0.06)	0.32
Missouri	0.26	0.37 (0.35, 0.40)	–0.08 (–0.12, –0.03)	0.45
Tennessee	0.34	0.46 (0.42, 0.49)	0.04 (–0.02, 0.11)	0.42

NOTE: The 90 percent coverage intervals are in parentheses.

(0.34 percent), which was quite a bit higher than the state with the second highest average growth rate, Indiana (0.30 percent). At the other end, Mississippi had the lowest average growth rate (0.25 percent), which was not far from the performances of Missouri and Illinois (0.26 percent and 0.27 percent, respectively).

Because the Hamilton model allows states to switch between expansion and recession phases, these actual average growth rates can be broken down into their two component growth rates. The second and third columns of Table 1 provide the estimated average growth rates in expansion and recession for the seven District states. Of these states, Tennessee's average growth during expansion (0.46 percent) easily outpaced that of Illinois (0.41 percent) and Indiana (0.39 percent). Of the remaining states, Mississippi has the lowest average expansion growth rate (0.34 percent).

It is during their recession phases that the differences between states are most glaring. Three states—Arkansas, Mississippi, and Tennessee—have average recession growth rates that are positive, although the 90 percent coverage interval around the estimates includes negative numbers. At the other end are three states—Illinois, Indiana, and Missouri—whose average recession growth rates are well below zero. These are also the states that suffer the most during a month of recession in that they have the greatest differentials between their average expansion and recession growth rates. In each of these states, a month of recession tends to mean about one-half of a percentage point lower growth. Tennessee is not far behind these states in terms of the opportunity cost of a month in recession. Although Tennessee tends to experience positive growth during a month of recession, its average growth rate in expansion is so high that its average output lost during recession is also relatively high.

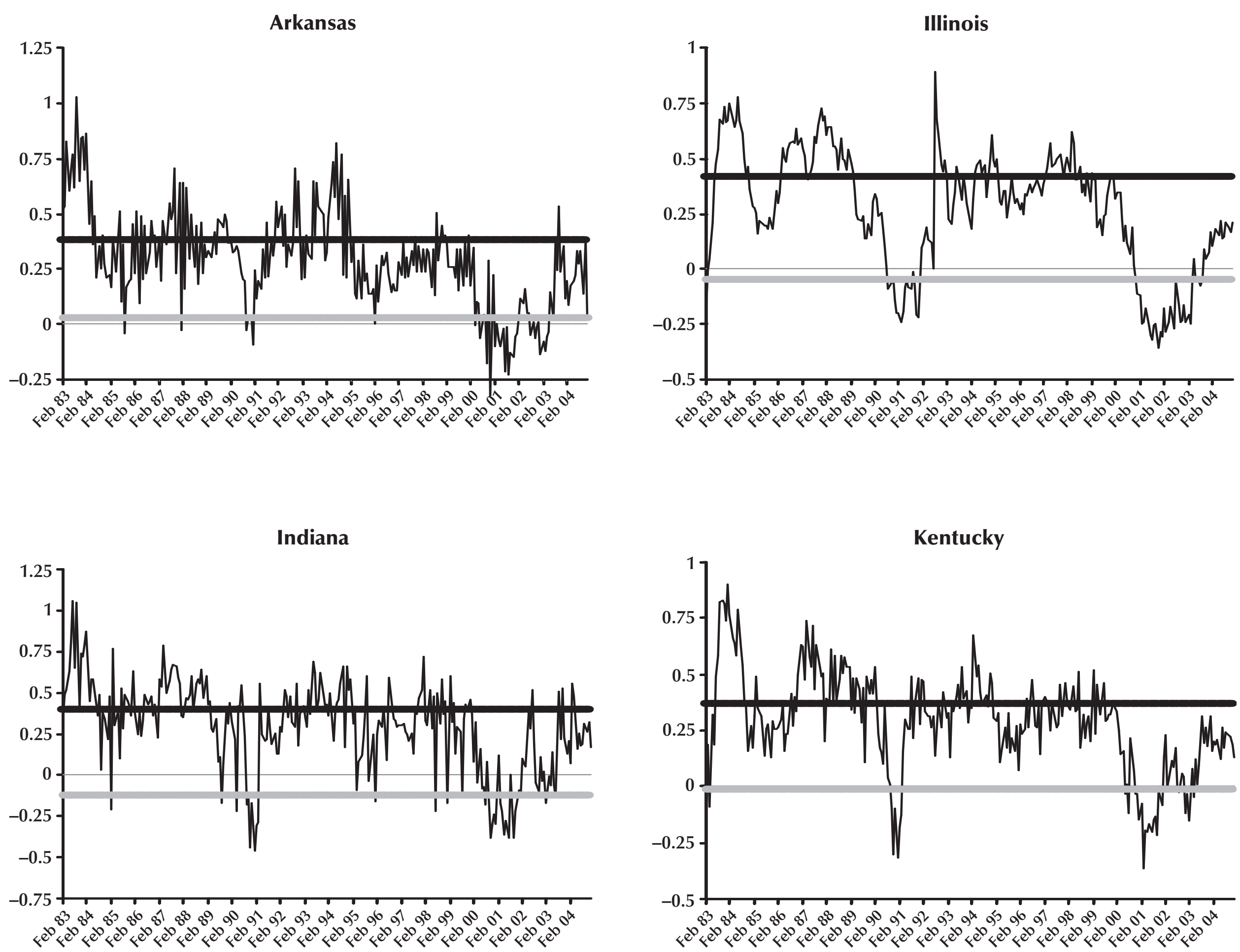
In trying to explain the differences in average growth rates across states, OPW found that the factors that might explain recession growth rates are not the same as those that explain expansion growth rates. Specifically, they found that demographic factors such as education and age distributions were related to expansion growth rates, but not to recession growth rates. On the other hand, they also found that states' shares of employment in manufacturing and construction and mining were related to recession growth rates, but not to expansion growth rates.

Pre-2001 Recession/Expansion Experiences

For reference, Figure 3 provides the actual monthly growth rates for the District states, along with the estimated average growth rates in the two business cycle phases. The Hamilton model also determines the probability that the data for a particular month indicates either a recession or expansion. It is not just data for the month in ques-

Figure 3

Actual and Average Growth Rates for District States (percent)



NOTE: Thick black lines are average expansion growth rates; thick gray lines are average recession growth rates.

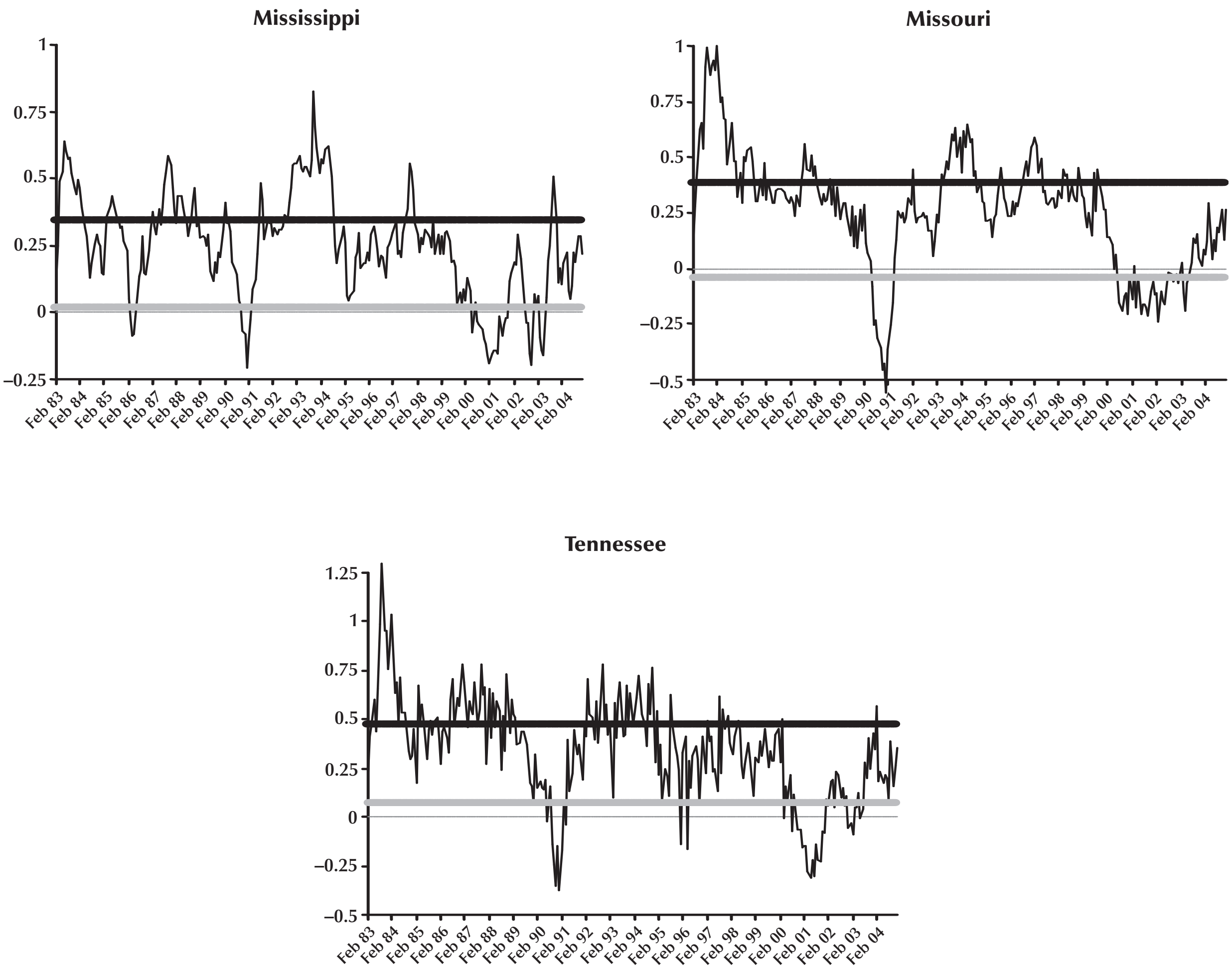
tion that matters, however. For a given month, the probability of being in recession also depends on the preceding and subsequent months.

The recession and expansion cycles for the seven District states are illustrated by Figure 4, which provides the probability of recession for each month of our sample period. In these charts, the light-shaded areas indicate the national labor-market recessions determined above and the dark-shaded areas indicate NBER recessions. The model does a good job of separating the data into two phases in that the probability of recession is usually close to either 1 or 0, although Indiana experienced several idiosyncratic spikes. On the whole, however, the model provides a clear picture of recent state-level recession/expansion experiences in the District.

Even though we deal with only a small subset of states, Figure 4 illustrates that some of the business cycle characteristics found by OPW for the 50 states apply to the 7 District states: Although state-level recessions tend to occur alongside national recessions, there have been occasions of state recessions that were independ-

Figure 3, cont'd

Actual and Average Growth Rates for District States



NOTE: Thick black lines are average expansion growth rates; thick gray lines are average recession growth rates.

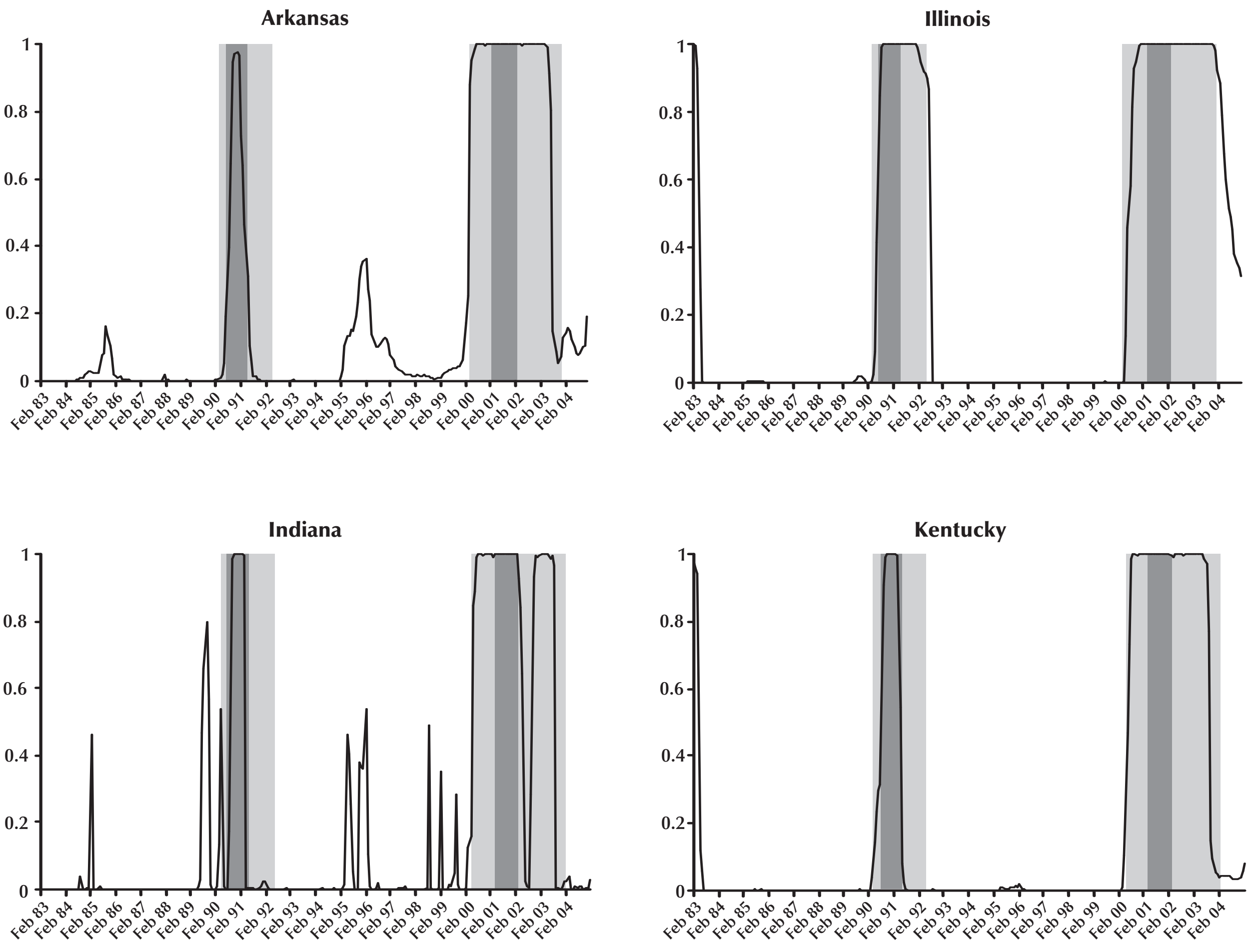
ent of national recession. In addition, there are significant state-level differences in the timing of recession episodes, relative both to each other and to the country as a whole.

For the most part, District states experienced labor-market recessions that were roughly in line with the NBER recession of 1990-91. In this regard, the District differed from other parts of the country, particularly the coasts, where labor-market recessions began much earlier and ended much later than the NBER recession. As described by OPW, there was a strong geographic pattern to the state-level labor-market recessions of the period: States in the Northeast and Far West switched into labor-market recessions up to 2 years before the start of the NBER recession. Recession spread from the coasts into the interior of the country and receded back to the coasts, ending for some Eastern and Western states more than 2 years after the end of the NBER recession.

There were interesting intra-District differences in the period. Of the District states, Illinois stands out in that its labor-market recession was very similar in timing to the national labor-market

Figure 4

State Recession Probabilities



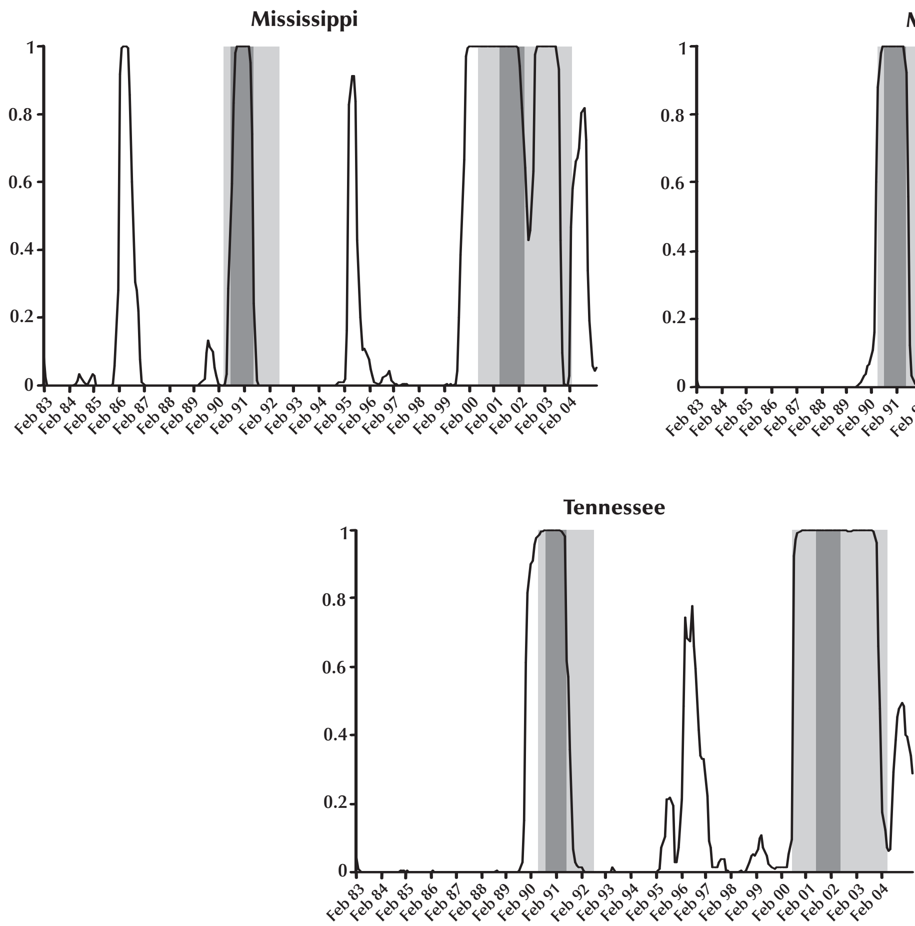
NOTE: Dark-shaded areas indicate NBER recessions; light-shaded areas indicate national employment recessions.

recession. Less glaring than the difference between the experience of Illinois and those of the rest of the states is that labor markets in Tennessee and Missouri went into recession 10 and 4 months, respectively, before the start of the NBER recession. Finally, the labor-market recessions in Arkansas, Indiana, Kentucky, and Mississippi were somewhat shorter than the NBER recession. It should also be noted that Indiana experienced a brief labor-market recession in 1989 that was not experienced in the rest of the District. Although some coastal states were in recession at this time, it is likely that Indiana's recession was idiosyncratic and unrelated to the recessions experienced in other parts of the country.

Although there were only two national recessions during our sample period, OPW found that there were two periods during which significant numbers of states went into recession while the national economy remained in expansion. The first such period was in 1985-86, when, following simultaneous downturns in the petroleum and agricultural sectors, nearly every state geographically between Idaho and Louisiana was in recession for at least one quarter. As Figure 4 shows, Mississippi was the only District state to have

Figure 4, cont'd

State Recession Probabilities



NOTE: Dark-shaded areas indicate NBER recessions; light-shaded areas indicate national employment recessions.

experienced this non-national recession, although Arkansas and Indiana experienced enough of a slowdown for their probabilities of recession to blip upward.

The second period during which states experienced labor-market recessions while the national economy was in expansion was in 1995. OPW found that several states—although no District states—experienced recessions beginning in 1995 that lasted between one and five quarters. As shown in Figure 4, we find that two District states—Mississippi and Tennessee—switched into recession during 1995 and that Arkansas and Indiana saw their probabilities of recession rise during the year without becoming high enough to indicate an actual recession. The difference between our results and those of OPW with regard to District states is likely due to our sample period, which, as mentioned above, was chosen to account for possible structural breaks during the 1980s.

THE 2001 RECESSION

As OPW showed, many state-level labor markets were not in sync with the NBER recession of

2001. States in the Mississippi and Ohio valleys, the Southeast, and the Northwest had switched into labor-market recessions in 2000. Also, many states located geographically between Montana and Texas had not switched into labor-market recession until mid-2001. Further, OPW found that most state-level labor-market recessions continued past the end of the NBER recession. Because their sample period ended in mid-2002, however, OPW was unable to provide a complete picture of the labor-market recessions at the state level. In this paper, with revised data through the end of 2004 in hand, we are able to analyze the entire labor-market recession experience of the states of the Eighth District during the period surrounding the 2001 NBER recession. Also, because of the differences in our sample period, we find presently that the labor-market recessions in District states began much earlier than had been documented by OPW.

Although this was not the case with the previous (1990-91) recession, the states of the Eighth District had labor-market recessions in and around 2001 that were much more in line with the national labor-market recession than with the NBER recession. Like the national labor market, the seven state labor markets went into recession well before the start of the NBER recession and into expansion long after the end of the NBER recession. As shown in Figure 5, which provides a close-up view of District states between 1999 and 2004, there were interesting differences within the District and between District states and the country as a whole. In the figure, a solid box (■) indicates that a state's labor market was in recession during that month. The dark-shaded area shows the 8-month-long NBER recession, and the light-shaded area shows the 42-month-long national labor-market recession. The cross-state comparisons are summarized in Table 2, which compares the timing of each state's labor-market recession with that of the national labor-market recession.

Arkansas

Arkansas's labor market switched into recession in April 2000, 2 months before the national labor market made the same switch, and a year before the start of the NBER recession. As one can see from Figure 3, Arkansas's growth dropped off dramatically during this month and remained low long afterward. Throughout the first half of 2003, growth began to rise until the labor market switched into expansion in August 2003, 4 months before the national labor market switched into its expansion phase. Overall, then, Arkansas's labor-market recession, which lasted 40 months, was 4 months shorter than the national labor-market recession. Among District states, only Indiana spent fewer months in recession.

Illinois

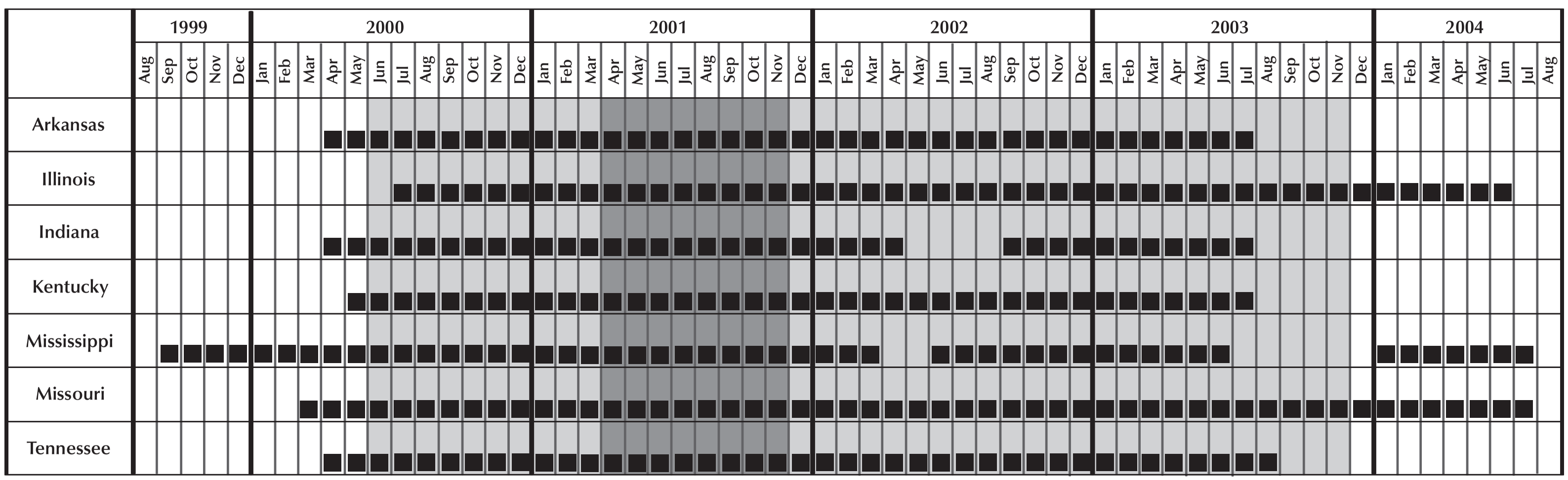
Illinois's labor market switched into recession in July 2000, 1 month after the national labor market did and 9 months before the start of the NBER recession. Of the District states, Illinois was the last to switch into recession. Figure 3 shows the clear drop-off in growth that occurred at that time and which continued into 2001 and beyond. It was only in late 2003 that growth became positive, although it was not until July 2004 that it became persistently high enough that the probability of recession went below the 0.6 threshold, thereby signaling the end of the state's labor-market recession. Note that even by the end of 2004, the expansion in Illinois was not as clear as in other states. This lack of clarity can be seen in Figure 3: Although growth had been persistently positive for several months, it was still languishing between the expansion and recession growth rates. Other states had made up more of the growth gap and had, therefore, seen a clearer signal that expansion was ongoing. In total, Illinois spent 49 months in labor-market recession during the period surrounding the 2001 NBER recession.

Indiana

Between the beginning of 2000 and the end of 2004, Indiana experienced two distinct labor-market recessions interrupted by a brief expansion. Indiana's first recession began in April 2000, 2 months before the start of the national labor-market recession. The state switched back into expansion in May 2002, 5 months into the NBER expansion. As illustrated in Figure 3, growth in

Figure 5

The 2001 Recession and the Eighth District States



NOTE: The solid boxes indicate the state's labor market was in recession. The dark-shaded area indicates the NBER recession; the light-shaded area indicates the national employment recession.

Table 2

Summary of Labor-Market Recessions Surrounding the 2001 NBER Recession

	Start of first recession period (U.S. = 0)	End of last recession period (U.S. = 0)	Number of recession periods (U.S. = 1)	Total months of recession (U.S. = 42)	Percent SCI reduction
Arkansas	–2	–4	1	40	–15.2
Illinois	+1	+7	1	49	–22.9
Indiana	–2	–4	2	36	–17.8
Kentucky	–1	–4	1	39	–16.5
Mississippi	–9	+8	3	51	–16.3
Missouri	–3	+8	1	53	–22.2
Tennessee	–2	–3	1	41	–22.0

Indiana again rose to above the average expansion growth rate and remained high until August 2002. By September 2002, however, the Indiana labor market had returned to recession, where it remained until August 2003, the first month of an expansion that it has so far maintained. Although Indiana experienced a "double-dip" labor-market recession, the state spent fewer months in recession (36) than any other District state.

Kentucky

Kentucky's labor-market recession began in May 2000, 1 month before the start of the national labor-market recession and 11 months before the NBER recession. Figure 3 illustrates the obvious drop-off in growth that signaled the switch into recession. As with the other states, growth rose during 2003, sustaining a level close to the average expansion growth rate by the middle of the year. The final month of Kentucky's single recession episode was July 2003, 4 months before the final month of the national labor-market recession and more than a year and a half after the end of the NBER recession. In total, Kentucky spent 39 months in a labor-market recession.

Mississippi

Mississippi was the first of any District state to switch into labor-market recession and was the last to enter a sustained period of expansion. Further, it experienced a "triple dip" in that it saw two short periods of recovery between three periods of recession. The first recession period began in September 1999, 9 months before the start of the national labor-market recession and 19 months before the start of the NBER recession. The first recovery period was in April and May 2002; but, as shown in Figure 4, the model produces a relatively high probability of recession even during these months. The second recession phase lasted until June 2003, which was about the time that several other District states were switching into expansion. Mississippi's expansion lasted only 6 months, however, and the third period of labor-market recession began in January 2004, 2 months after the national labor-market recession ended. By July 2004, this final, 7-month-long recession period ended, 8 months after the end of the single national labor-market recession. Between September 1999 and July 2004, Mississippi's labor market was in recession for 51 months and in expansion for 8 months.

Missouri

Of the seven District states, Missouri spent the most months in recession (53) during the period surrounding the 2001 NBER recession. Whereas Mississippi's overall recession experience spanned a longer time period, Missouri's labor market was in recession for a single block of time that began in March 2000, more than a year before the start of the NBER recession and 3 months before the national labor-market recession. As

Figure 3 shows, Missouri's growth dropped off considerably at this time, providing a clear signal that its labor market had switched into recession. Growth remained consistently negative through the rest of 2000 and all of 2001 and 2002. Although it picked up somewhat during 2003, the model yields little doubt that the recession continued through to 2004. It wasn't until August 2004, 8 months into the national labor-market expansion, that growth was high and persistent enough to signal the start of the labor-market expansion.

Tennessee

Tennessee's labor-market recession began in April 2000, 2 months before the start of the national labor-market recession (The drop-off in growth at this time is clear from Figure 3.) Its recession ended shortly after those of Arkansas, Indiana, and Kentucky, and 3 months earlier than the national labor market. In total, Tennessee's labor market was in recession for 41 months. Within this single recession phase, Tennessee's growth was relatively turbulent, falling well below its average recession growth rate for all of 2001. Recovery was strong in 2003, however, and growth had risen above the average expansion growth rate by the beginning of 2004. More troubling for Tennessee was the rocky performance during the second half of 2004, when growth fell and the probability of recession rose, as shown in Figures 3 and 4. At that time, however, growth had not fallen persistently close enough to the average recession growth rate to have signaled a recession.

CONCLUSIONS

Typically, District states experienced labor-market recessions that were roughly in line with the NBER recession of 1990-91. This is in contrast with other parts of the country where labor-market recessions began much earlier and ended much later than the NBER recession. Illinois differed from the rest of the District states in that its labor-market recession was very similar in timing to the national labor-market recession. In addition, the labor-market recessions in Arkansas, Indiana, Kentucky, and Mississippi were somewhat shorter than the NBER recession was.

During the period surrounding the 2001 NBER recession, the labor markets of four District states—Arkansas, Indiana, Kentucky, and Tennessee—spent somewhat less time in recession than did the national labor market. Each of these states went into labor-market recession a month or two before the country as a whole, while entering expansion 3 to 5 months earlier than the country as a whole. On the other hand, the labor markets of three District states—Illinois, Mississippi, and Missouri—were in recession for more time than the national labor market was. All three switched into recession earlier than the country as a whole: For Mississippi, the switch occurred 10 months earlier. Sustained labor-market expansion didn't begin in these three states until 7 or 8 months after it did for the country as a whole.

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Cyclical Patterns and Structural Changes in the Louisville Area Economy Since 1990

Paul Coomes and Nan-Ting Chou

In this paper, the authors examine several data sets to better understand the growth patterns in the Louisville area economy since 1990. They find that the regional economy has closely tracked the national economy, in terms of growth in jobs, payroll, and housing. However, changes in job-based location quotients suggest that the structure of the Louisville economy has actually diverged from the national economy over the period. Manufacturing overall has declined in the Louisville area as it has nationally, but the local motor vehicle and related parts manufacturing subsector has doubled in relative importance due to the southward movement of the automobile and truck industries. Also rising in relative importance are Louisville's distribution, recreation, and health services industries. The authors examine these growth patterns and structural changes and also investigate the false signals about the local economy sent by the U.S. Bureau of Labor Statistics data before and during the recession of 2001.

Federal Reserve Bank of St. Louis *Regional Economic Development*, 2005, *1*(1), pp. 17-29.

If you think of Fed Chairman Greenspan as having access to a fire hydrant of data on the U.S. economy, then think of regional economists as having access to a slow and leaky garden hose. Data arrive in dribbles long after events transpire, and there is never enough. For example, it often takes two years for estimates of monthly job growth to settle down, and revisions are particularly large around cyclical turning points in the economy. Current labor force estimates, except for heavily populated metro areas, are subject to large measurement errors. And there are generally no data at all on such key economic variables as regional investment, prices, industrial output, retail sales, trade, wealth, and consumer spending. Hence, analysts are left to sort out the structural changes and cyclical behavior of their regional economy retrospectively using the best data available, however skimpy and noisy.

In this paper, we examine several data sets to better understand the growth patterns in the Louisville area economy since 1990. We find that the regional economy has closely tracked the national economy, in terms of growth in jobs, payroll, and housing. Overall manufacturing employment has declined, and distribution, recreation, and health services jobs have risen in importance. Job growth has recently picked up in Louisville, but, similar to the nation as a whole, nonagricultural wage and salary employment has not yet reached its pre-recession peak (Lloyd and Meuller, 2005). Wages and salaries have also grown in synch with payrolls nationally. Similarly, new home construction was relatively strong throughout the recession and recovery, both locally and nationally. However, changes in industrial location quotients suggest that the structure of the Louisville economy has actually diverged somewhat from the national economy. We examine these growth patterns and structural changes and

Paul Coomes is professor of economics at the University of Louisville and the National City research fellow. Nan-Ting Chou is an associate professor of economics at the University of Louisville. The authors thank Barry Kornstein and Margaret Maginnis for data, mapping, and analytical support.

Table 1

Growth in Macro Variables, the Louisville MSA vs. the U.S.

	1990	2004	Growth (percent)
Jobs, nonagricultural wage and salary (000)			
Louisville MSA	512	599	17.0
U.S.	109,487	131,480	20.1
Population, July 1			
Louisville MSA	1,058,425	1,200,063	13.4
U.S.	249,622,814	293,871,612	17.7
Single-family building permits/starts			
Louisville MSA—permits	3,680	6,184	68.0
U.S.—starts (000)	895	1,613	80.2

NOTE: The Louisville MSA refers to the new 13-county definition, except for building permits, which are based on the pre-2004 7-county definition.

SOURCE: Job data are from the U.S. Bureau of Labor Statistics. Population data are from the U.S. Bureau of Economic Analysis, with estimates for 2004 based on extrapolation of 2003 data. Building permit data are from the U.S. Census Bureau.

also investigate the false signals about the local economy sent by U.S. Bureau of Labor Statistics (BLS) data before and during the recession of 2001.

DEFINITION AND GEOGRAPHIC SCOPE OF LOUISVILLE ECONOMY

The Louisville metropolitan statistical area (MSA) was redefined in 2003 to include 13 counties in Kentucky and southern Indiana. The total MSA population is about 1.2 million. Commuting patterns revealed by the 2000 Census led to the geographic expansion of the MSA, from its former 7-county definition.

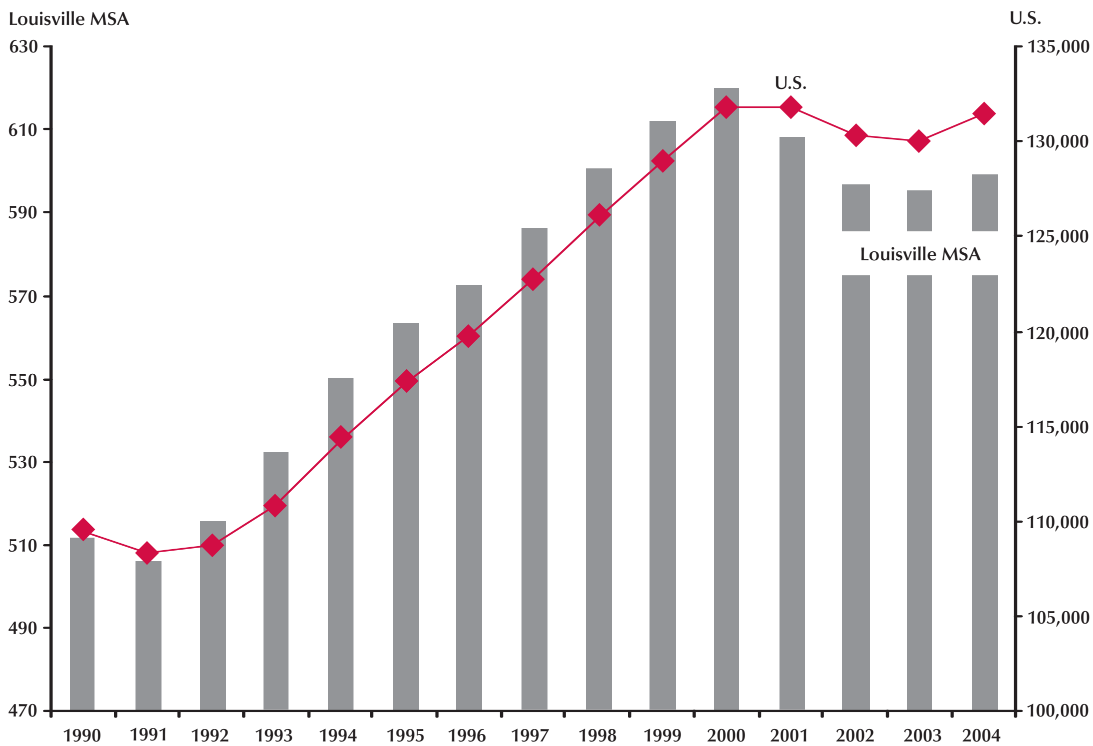
The Louisville MSA, as well as the newly defined Elizabethtown MSA, is part of the wider 25-county Louisville economic area. The economic area classification was developed by the U.S. Bureau of Economic Analysis and assigns all U.S. counties to some regional economy. This broader definition is very useful in analyzing the markets for labor, major retail purchases, television and print media, air transportation, higher education, and major medical and professional services. In fact, in previous studies we have found the MSA geography to be too small to account for the labor force growth occurring on a place-of-work basis in the MSA (Coomes et al., 2000). Low interest rates, combined with the relatively recent real estate developments around outlying interstate highway interchanges, the raising of the speed limit in the early 1990s, low real gas prices, improved automobile efficiency, and the demand for inexpensive modern homes have all caused people to live further and further from their workplaces. In terms of population, the fastest growing counties in the Louisville economic area have been in the first and second rings around the central county—Jefferson County, Kentucky. Under the broader economic area definition, the Louisville economic area is bounded by the Indianapolis, Cincinnati, Lexington, Nashville, and Evansville economic areas. However, because much more economic data are available for MSAs than for economic areas, the rest of the paper focuses on the economy of the Louisville MSA.

AGGREGATE ECONOMIC MEASURES

The economy of the Louisville MSA was relatively strong in the 1990s, particularly compared with its growth during the previous two

Figure 1

Total Jobs, the Louisville MSA vs. the U.S.



NOTE: Jobs are in thousands. Louisville MSA refers to the 13-county definition, as of 2004.

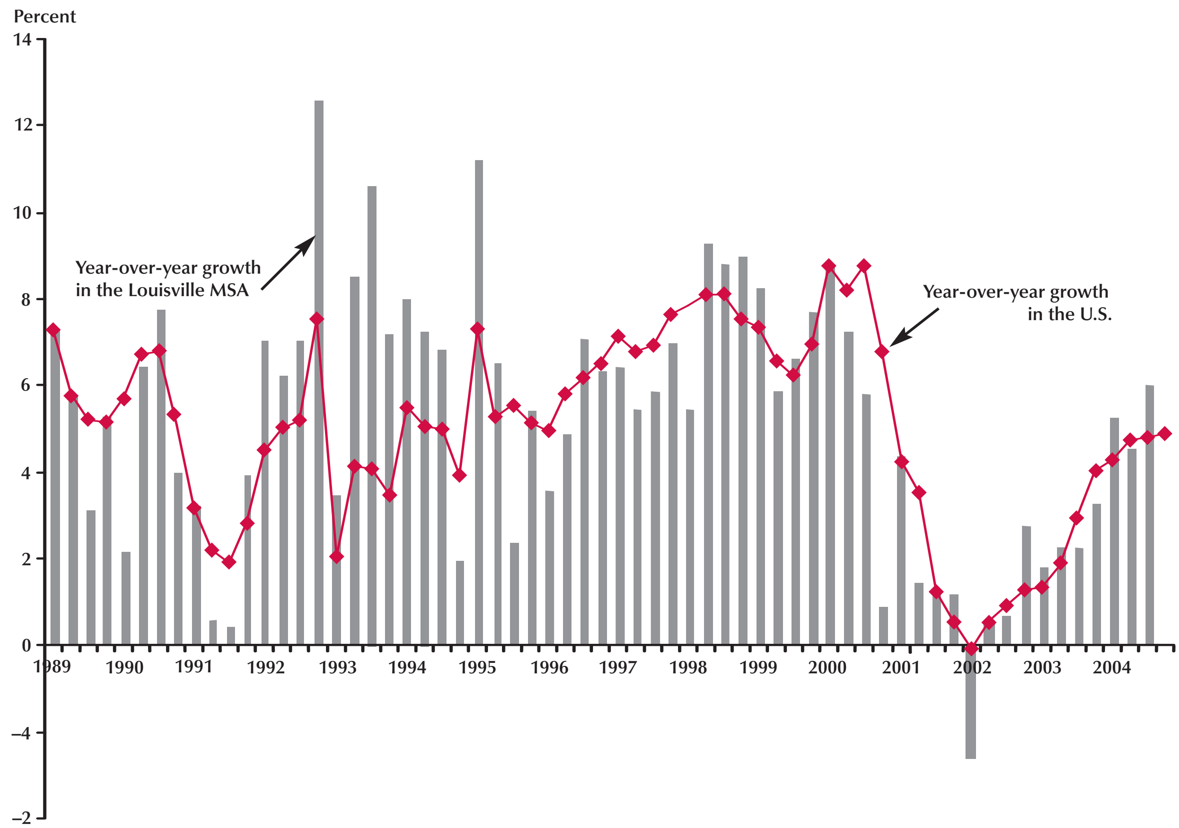
SOURCE: BLS, Current Employment Survey; nonagricultural wage and salary jobs only, in thousands.

decades. The continued expansion of the international air freight hub of United Parcel Service was a key driver, and the region added thousands of jobs in distribution, warehousing, logistics, and related transportation industries. The MSA also benefited from the southward movement of the U.S. automobile industry, seeing job growth at its two large Ford assembly plants and at many parts plants around the region. This good job growth induced an acceleration of population growth. The population and income growth, combined with interest rates at generational lows, stimulated the housing market. Of course, the 1990s were a bullish time for the national economy as well. Yet, the Louisville MSA's growth since 1990 was slightly below the national average for key aggregate economic measures. See Table 1.

The 13-county Louisville MSA gained 87,000 jobs between 1990 and 2004, with all the net growth occurring between 1992 and 2000. Indeed, according to the BLS, the economy of the Louisville MSA now supports 20,000 fewer jobs than it did at the end of the previous decade. Over the entire 14-year period, Louisville's net growth was 17 percent, compared with national growth of 20 percent. Figure 1 shows the nearly contemporaneous growth in jobs for the Louisville MSA and the United States over the expansion period. However, the MSA apparently suffered a greater percentage loss of jobs during the recession than did the United States as a whole. The slight uptick in jobs shown for the Louisville MSA in 2004 is based on preliminary estimates, and we have learned to be suspicious about early job

Figure 2

Quarterly Growth in Wages and Salaries Paid, the Louisville MSA vs. the U.S.



SOURCE: BLS, wages and salaries covered by unemployment insurance; both series are seasonally adjusted.

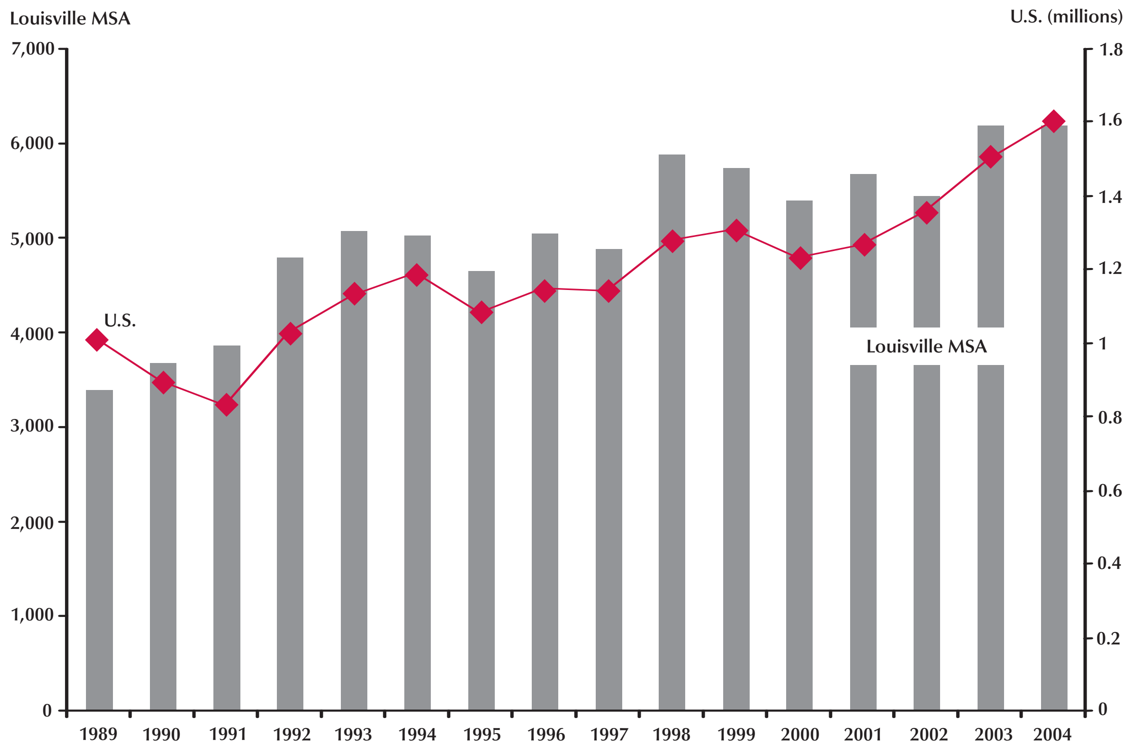
estimates for the local economy. Revisions to the MSA's job data for 2004 (perhaps upward) are likely to be much greater than revisions for the nation as a whole. We discuss revisions to job data later in the paper.

Conventional wisdom among economic development and civic leaders in the Louisville MSA is that the economy lags the national economy. If this were ever true, it is certainly no longer. Consider data on payroll growth at a quarterly frequency. The BLS compiles data on wages and salaries covered by the unemployment insurance system. We compare the year-over-year quarterly growth rates of payrolls in the Louisville MSA (old 7-county definition) with those of the nation, as shown in Figure 2. The data are seasonally adjusted so that the underlying pattern is easier to detect. It is clear from the chart that the growth in wages and salaries paid in the Louisville MSA tracks closely with that of the nation, especially over the past few years. This makes us wonder if there has been a structural convergence, wherein the Louisville MSA has become statistically more like the nation.

Like the nation as a whole, the Louisville MSA's housing market continued to thrive throughout the previous recession. Figure 3 illustrates growth in single-family home building permits issued in the Louisville MSA versus housing starts in the United States over the past 15 years. The reasons for the growth are now well-known, though few forecast such a strong housing sector

Figure 3

Louisville MSA Single-Family Home Building Permits vs. U.S. Housing Starts



SOURCE: U.S. Census Bureau.

during the cyclical downturn. The most common explanations include a mixture of very low interest rates for both developers and home buyers, an expanding base of retired persons seeking new homes, declining household size due to divorces and double income earners deciding to have fewer children, and the low expected return on alternative forms of investment like equities and money-market instruments. Interestingly, the growth in single-family homes nationally exceeded the growth in the Louisville MSA by about the same ratio as that for jobs and population.

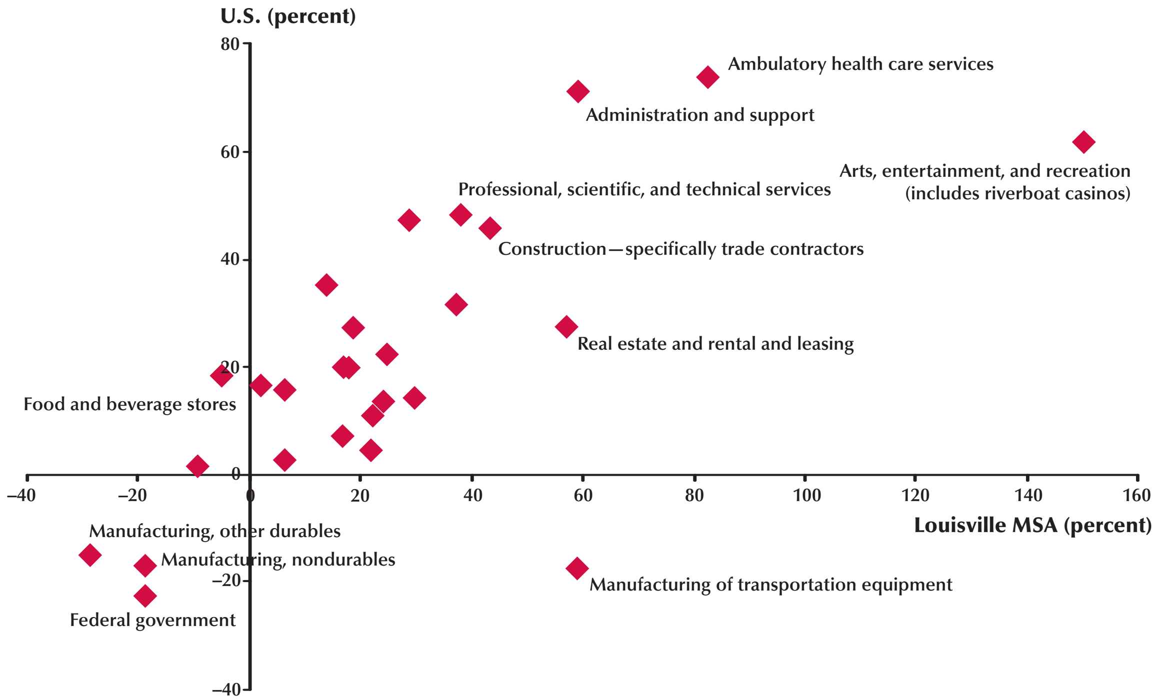
JOB GROWTH BY INDUSTRY SINCE 1990

At the aggregate economic level, we see a remarkable similarity between the growth in the Louisville MSA and the United States over the past 14 years. We now investigate this further, using job data on 26 industries. We calculate location quotients, using detailed job data to learn which local industries have gained and lost in importance compared with the same measures nationally. We find that the variation in location quotients actually increased since 1990, suggesting that the structure of the Louisville economy has actually diverged somewhat from the national economy. Given the very similar growth paths revealed by aggregate data on jobs and payrolls, these structural changes must have offsetting effects on employment and payrolls at the aggregate level.

In-depth analyses of regional economies are hampered by the break in most time series due to the introduction of the North American Industrial

Figure 4

Job Growth Rates by Industry, the Louisville MSA vs. the U.S., 1990 to 2004



SOURCE: BLS.

Classification System (NAICS), which has replaced the old Standard Industrial Classification (SIC) system. The NAICS conventions have cleaned up past misclassifications, such as treating corner bakeries and copying shops as manufacturers. However, because of the lack of sufficiently detailed historical data, NAICS-based estimates have been published only retroactively to 1990 for MSA-level jobs and to 2001 for earnings by industry. The recent changes to MSA definitions, aligned with the commuting patterns data from the 2000 Census, also create some comparability problems in public data sets. Hence, in this paper, we blend data from different sources using different definitions, to tease out stories about the recent path of the Louisville economy.

Figure 4 provides a scatter plot of job growth by industry for the Louisville MSA and the United States. These are the data, released in March 2005, that provide retroactive estimates using for the first time the NAICS, and which also correspond with the 2004 geographic redefinitions of MSAs. Appendix Table A1 provides reference data. We are limited in how much industrial detail we can analyze because the BLS provides job estimates for only 26 industries in Louisville. Yet, this is enough detail to reveal several interesting patterns. First, the growth rates of most local industries are in line with national rates. The slope of a regression line through the scatter plot is 0.98, and the correlation coefficient between the growth rates is 0.70. The two fastest growth industries nationally—ambulatory health care services and administrative support—were also among the fastest growing locally. Ambulatory health care primarily refers to outpatient care facilities and offices of physicians, dentists, optometrists, psychologists, physical therapists, and labs. The administrative and support sector refers to a large collection of industries that do "office administration, hiring and placing of personnel, document preparation and similar clerical services, solicitation, collection, security and surveillance services, cleaning, and waste disposal services,"[1] services that are sometimes performed in-house by companies and households. The three industrial categories that lost jobs nationally—federal government, nondurable manufacturing, and durable manufacturing (less transportation equipment)—lost jobs locally as well. Louisville also lost jobs in the food and beverage retail-store category, though this industry posted a slight gain nationally. Retail, the second fastest-growing sector (after services) in terms of jobs over the past two decades, is evidently undergoing a technological revolution, because there has been no net job growth for the past few years even though retail sales continue to advance.

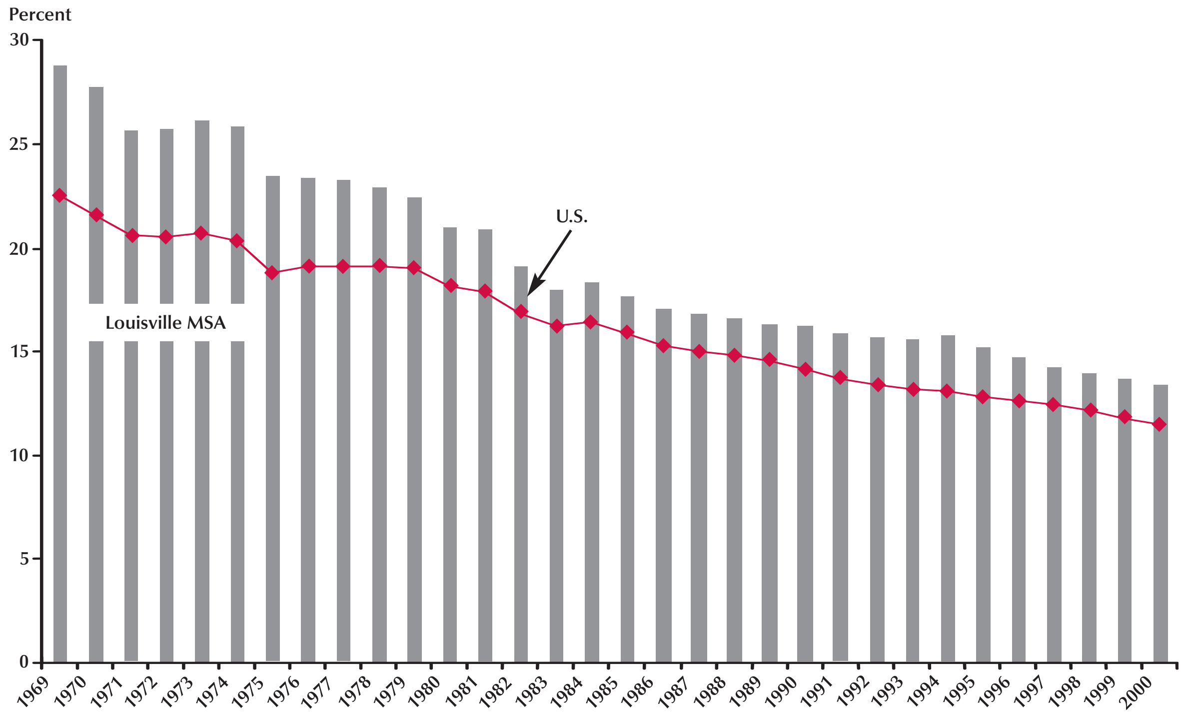
One expects to see more variation locally than at the national level because the national economy has a much greater number and variety of firms in each industry, which tends to generate offsetting effects. No MSA economy perfectly mirrors the national economy, because the particular climate, history, industrial structure, demographic characteristics, and policies of a place will favor certain industries and disfavor others relative to the national average. In this regard, there are two obvious outliers in the Louisville MSA data. The MSA posted much higher job growth than the nation in the arts, entertainment, and recreation industry and in the manufacturing of transportation equipment industry. Although published BLS data do not reveal this, we know from other information that the growth in arts, entertainment, and recreation was really due to the opening of a large riverboat casino in Harrison County, Indiana, part of the Louisville MSA, and a few miles west of downtown Louisville. The Caesar's casino opened in late 1998 and its hotel in 2001; combined they have about 2,000 employees, over one-fifth of the MSA's jobs in the entire arts, entertainment, and recreation industry. Louisville's new relative specialization in this industry shows up in the change in location quotients over the period. In 1990, the industry's location quotient was 0.72 (net importer). By 2004, the location quotient was 1.14 (net exporter).

The other local industry to significantly outperform the nation was the manufacturing of transportation equipment. The Louisville MSA has benefited from its central location along the north-south automobile and truck production corridor. The area has two major Ford assembly plants, with total employment of about 9,000. One makes the popular Ford Explorer and other SUVs; the other makes heavy-duty trucks, F-250s through F-550s. Although sales have been soft during the past few months, over the past few years the plants have gained business as Ford has shifted production from Cleveland, St. Louis, and other markets to Louisville. The area also has many auto-related parts plants, including a ThyssenKrupp Budd body-stamping plant in Shelbyville that employs 1,100 and a Tower Automotive frame plant in Corydon that employs 800. The transportation equipment manufacturing sector as a whole grew in employment locally from 11,000 to 17,000 jobs between 1990 and 2004, with only a slight dip during the previous recession. Note that this strong growth occurred against a backdrop of declining national employment in the industry. Kentucky in general and the Louisville area in particular now boast the lowest industrial electricity rates in the country, thanks to the California-driven spikes in power prices in the northwestern United States in the late 1990s and early part of this decade. Perhaps more important than energy costs, though, is the location advantage. Our deceased colleague Mark Berger of the University of Kentucky used geographic information system tools to calculate the geographic center of the United States population east of the Rocky Mountains. He found that it is near Danville, Kentucky, just south of Louisville. This means that if one needs to ship heavy, expensive products to major consumer markets, there is no better way to minimize transportation costs than to locate your plant in central Kentucky. Two major north-south interstates, I-65 and I-75, pass through this area and connect the traditional auto-production heartland of Michigan, Ohio, and Indiana to the emerging areas of Kentucky, Tennessee, and Alabama. Indeed, half of all manufacturing jobs created in Kentucky during the past decade were in a handful of counties in the center of the state.

[1] See www.census.gov/epcd/naics02/def/NDEF56.HTM#N56.

Figure 5

Manufacturing's Share of Total Jobs, the Louisville MSA vs. the U.S.



NOTE: Manufacturing industries are classified on the old SIC basis. Louisville MSA refers to the 13-county definition. Jobs include both labor and proprietors.

SOURCE: U.S. Bureau of Economic Analysis.

We previously noted here the decreasing reliance on manufacturing employment in the Louisville MSA, commensurate with a similar decline nationally. Figure 5 shows the declining share of manufacturing jobs over the last one-third of the century. Manufacturing's share of all jobs in the Louisville MSA fell from 29 to 13 percent, while the share nationally fell from to 23 to 12 percent. This longer time series is available only on an SIC basis and only through 2000; nevertheless, the convergence of Louisville to the national average is evident in the trend for manufacturing, at least at this aggregate level. This observation and the concurrence of growth patterns for total jobs and total payrolls led us to the investigation of jobs by industry.

While in general the same sectors grew (shrank) in Louisville as in the United States, we do not find a movement toward the U.S. distribution of jobs across industries. Of the 26 industries examined, 11 moved closer to the national share and 15 moved away from the national share. Of course, these share calculations are unweighted, so that industries with relatively few employees are counted the same as those with many employees.

We next calculated location quotients for the 26 local industries, for both 1990 and 2004. The standard deviation of the location quotients rose from 0.21 in 1990 to 0.31 in 2004. Thus, the descriptive evidence suggests that the industrial structure of the Louisville MSA has diverged statistically from that of the nation. Given that the aggregate measures of job and payroll growth in the MSA track the national paths so closely, this suggests that although the industrial structure of the MSA changed, the impacts on overall employment and payrolls tended to cancel each other out. That is, the gains in jobs and payrolls for the MSA relative to those for the nation in sectors such as transportation equipment manufacturing, wholesale trade, and entertainment were offset by the decreasing competitiveness of sectors such as information, professional and business services, retail trade, other educational and health services, and federal and state governments.

MEASUREMENT ERROR: JOBS AND UNEMPLOYMENT

Finally, we examine the job-data revisions that occurred during and after the 2001 recession, to learn more about the false signals that often occur in regional economies during turning point. We also look at local unemployment rates published at the time of the previous decennial census and see how they had exaggerated the degree of tightness in the local labor market.

The previous recession has been officially dated to begin in March 2001 and to end in November 2001. This is roughly contemporaneous with the pattern of growth in the national payroll series for nonagricultural wage and salary jobs, as generated by the Current Employment Statistics program of the BLS (2004). Payroll job growth peaked in March 2001 at 132.5 million (seasonally adjusted). Payroll job growth troughed in February of 2002 at 130.4 million. These data, because they are based on such a large sample of establishments from all over the country, are not subject to great revision when the BLS benchmarks to the unemployment insurance data (administered at the state level). The BLS reports that the average benchmark revision for total nonfarm employment over the past 10 years was just 0.3 percent (BLS, 2004). However, the revisions can be great at the local level, where monthly survey sample sizes are smaller, and particularly large at turning points. The signals during a recession can be very noisy, and unfortunately this was true for the Louisville MSA in 2001. To highlight this, we have organized data for the third quarter of 2001 and compared the preliminary BLS estimates with the final benchmarked estimates.

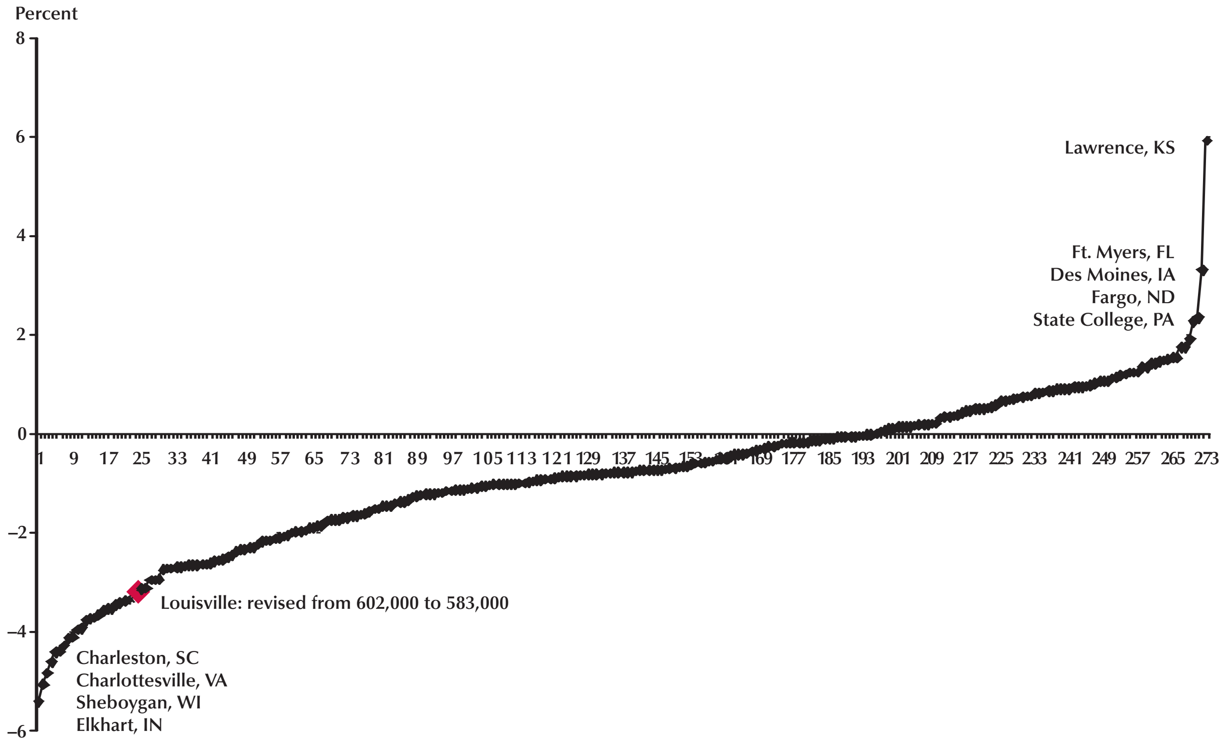
In late 2001, the BLS published its initial estimates for the July to September quarter, showing an average of 602,000 total nonagricultural wage and salary jobs in the seven-county Louisville MSA. This implied that, in the heart of the recession, the economy was still creating jobs—a total of 8,000 more than in the third quarter of 2000. The BLS ultimately revised the job estimate for that quarter to be 582,000, representing a net loss of 12,000 jobs from the prior year. The "interbenchmark" revisions of monthly employment, covering the period April 2001–February 2002, were released by the BLS in February 2003. By that time, the recession was over and the recovery phase was a year old. Yet, throughout the recession, local media used the BLS data to report that the Louisville MSA continued to add jobs, confusing the business community and government leaders who were seeing clear signs of a downturn in their enterprises.

Unfortunately, early job estimates for the Louisville MSA were among the noisiest in the nation. Revisions of job estimates for all 273 MSAs are shown in percentage terms in Figure 6. At –3.2 percent, Louisville's downward revisions were the twenty-fourth greatest among the MSAs. The revisions ranged from –5.4 percent (Elkhart, Indiana) to 5.9 percent (Lawrence, Kansas). It is clear, however, that most (197) MSAs saw a downward revision. The median revision among the metros was –0.8 percent, and the average revision was –0.9 percent.

Revisions varied widely by industry. Table 2 shows both the mean and median revisions for finance, insurance, and real estate (FIRE) and government were positive across the 273 MSAs. As in previous recessions, the construction industry was subject to the highest revision among MSAs (Coomes, 1992). Given the strong growth in single-family housing during this period, we presume the downward revision was due to an overestimate of nonresidential construction growth. The wholesale and retail trade sector, particularly retail trade, is very difficult to track during turning points in the economy. There are many small establishments, but they account for a large portion of total employment in the economy. The BLS surveys a very small fraction of

Figure 6

Revisions to Job Estimates for All 273 U.S. MSAs, 2001:Q3



SOURCE: BLS.

Table 2

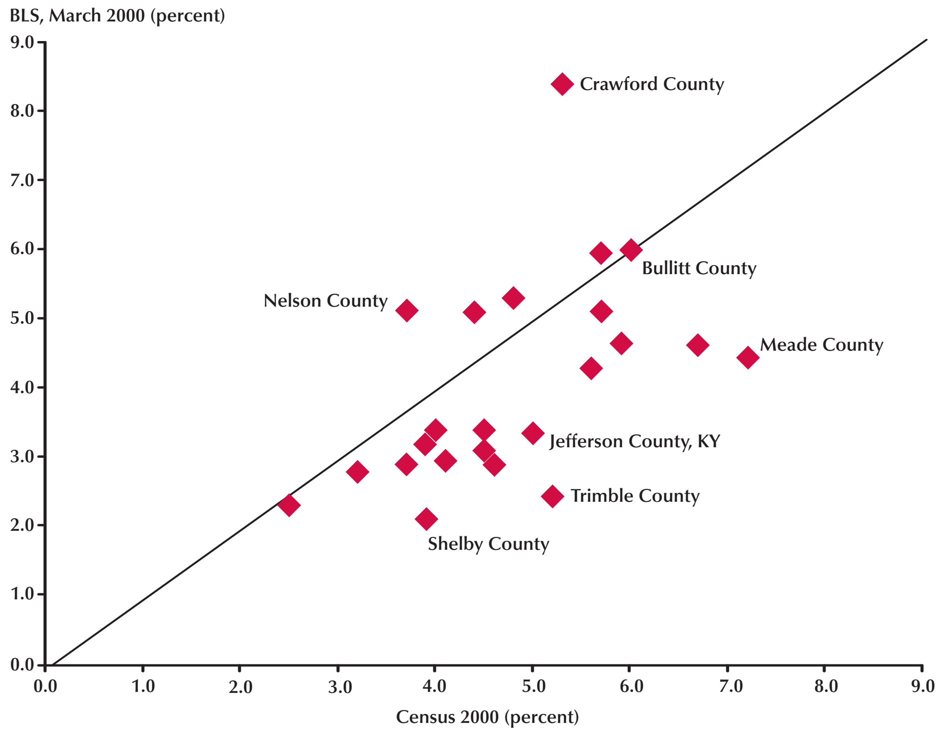
Revisions to Job Estimates, 2001:Q3

		273 MSAs	
	Louisville MSA (percent)	Mean (percent)	Median (percent)
Construction and mining	–1.02	–1.85	–1.92
Manufacturing	–0.20	–0.95	–1.27
Transportation, communications, and utilities	–0.99	–0.44	–0.16
Wholesale and retail trade	–6.98	–1.40	–1.50
Finance, insurance, and real estate	–4.34	0.84	0.41
Services	–3.05	–0.74	–0.72
Government	–1.09	0.31	0.21
Total nonagricultural	–3.19	–0.85	–0.78

SOURCE: Raw data are from the BLS, SIC basis, pre-2003 MSA definitions. Calculations are by the authors.

Figure 7

Unemployment Rate Estimates: Louisville Economic Area, BLS vs. Census 2000



NOTE: For the total 23-county area, BLS (3.6%) vs. Census (4.8%).

retail establishments and uses the reported employment levels to infer employment for the entire industry. Retail has been undergoing great technology-induced structural change, and hence correcting for sample bias is difficult. Only during the annual benchmarking process, when the census counts of all establishments and employees are compared with survey-based estimates, does one reveal the extent of the bias. We suspect that during the previous downturn, firms accelerated their substitution of labor with information technology equipment and systems. Moreover, less-flexible firms went out of business and were replaced by more innovative and flexible firms. The former were represented in the Current Employment Statistics survey, the latter were not. If so, that would explain the upward bias in retail employment estimates during the recession.

For Louisville, the initial employment estimates were too high for every industry. The greatest error was in the wholesale and retail trade sector, consistent with the national pattern, though much larger. The large revision in the service industry is presumably due to the same factor as retail trade—the industry is characterized by many small establishments that collectively support 30 percent of all jobs. We do not have insights into the source of the large error in the estimate for jobs in Louisville's FIRE sector. In summary, when looking for signs of recessions and upturns, analysts need to be very careful when referring to initial job estimates for MSAs.

We turn now briefly to a discussion of measurement error for local unemployment estimates. Published unemployment rates for MSAs and counties are treated with more suspicion than the current job estimates just discussed, but newspapers continue to highlight local unemployment data. Economic development leaders and elected officials continue to cite these statistics as if they accurately reflect local labor market conditions. Perhaps this is because there are so few opportunities to compare the BLS-produced estimates to the results of a large survey of households. The decennial census presents such an opportunity. During the U.S. Census 2000, conducted in the spring of that year, one in six households was asked to fill out the long-form questionnaire that probes for socioeconomic details. The Census questionnaire has a script that conforms closely to the employment and unemployment concepts used by the BLS.

At the time, there was great concern locally about an emerging labor market shortage and how that would strangle the steady growth the region had been witnessing for the prior eight years. The BLS estimated that there were only 18,200 unemployed persons in the Louisville MSA during the spring of 2000, reflecting an unemployment rate of 3.2 percent. When the Census results were released late the next year, we learned that in fact there were 24,400 unemployed persons in that period, reflecting an unemployment rate of 4.6 percent. The BLS had underestimated the number of unemployed persons by 6,000 (and overestimated the number of employed persons by 42,000). In Jefferson County, by far the most populous county in the region, the BLS had estimated the unemployment rate to be 3.4 percent for March 2000, whereas the Census estimate for the same period was 5.0 percent. This represents a difference of 5,000 people looking for work. Figure 7 shows the measurement error for all the counties in the Louisville economic area, as defined at the time by the U.S. Bureau of Economic Analysis. The larger economic area is a better geographic scope for the regional labor market than the (former) 7-county MSA (Coomes et al., 2003). According to the Census, there were 7,200 more unemployed persons in the region than had been estimated by the BLS. This implies that there was actually plenty of labor capacity to support economic growth locally, rather than the labor shortage widely perceived at the time.

CONCLUSION

Over the past 15 years, the Louisville MSA economy has experienced a growth path similar to that of the United States as a whole. At the aggregate level, the data on employment, payrolls, population, and housing starts for the MSA show similar economic trends and fluctuations as those for the nation. However, despite the evident convergence at the aggregate level, an examination of detailed industrial developments suggests increasing divergence from the national economy. This is indicated by the increasing standard deviation of Louisville MSA location quotients, using data on jobs in 26 industries, over the past decade. Finally, local economic data are subject to larger revisions. Although Louisville's job growth during the recent recovery was not as strong as national growth, it is possible that future data revisions will narrow the measured difference.

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Appendix Table A1

Nonagricultural Wage and Salary Jobs, by Available NAICS Industry Detail, 1990-2004

	Growth in jobs (percent)		Share of total jobs, 1990 (percent)		Share of total jobs, 2004 (percent)		Louisville location quotients	
	Louisville MSA	U.S.	Louisville MSA	U.S.	Louisville MSA	U.S.	1990	2004
Total nonfarm	17.0	20.1	100.0	100.0	100.0	100.0		
Specialty trade contractors	43.2	45.8	3.0	2.8	3.7	3.4	1.09	1.10
Other construction and mining	21.6	4.5	1.9	2.7	2.0	2.4	0.69	0.83
Manufacturing of transportation equipment	58.9	–17.3	2.1	1.9	2.8	1.3	1.07	2.12
Manufacturing of other durable goods	–18.9	–16.8	7.9	7.9	5.5	5.4	1.01	1.01
Manufacturing of nondurable goods	–19.0	–22.3	7.1	6.4	4.9	4.1	1.12	1.20
Wholesale trade	16.5	7.3	5.0	4.8	4.9	4.3	1.03	1.15
Food and beverage stores	–9.6	1.7	2.4	2.5	1.9	2.1	0.96	0.88
General merchandise stores	23.9	13.7	2.1	2.3	2.3	2.2	0.93	1.04
Other retail trade	–5.1	18.5	8.4	7.2	6.8	7.1	1.16	0.95
Transportation and utilities	29.5	14.3	5.6	3.9	6.2	3.7	1.45	1.68
Information	1.9	16.7	2.1	2.5	1.8	2.4	0.84	0.75
Finance and insurance	17.5	19.8	5.3	4.5	5.3	4.5	1.16	1.16
Real estate and rental and leasing	57.1	27.6	1.0	1.5	1.3	1.6	0.64	0.81
Professional, scientific, and technical services	37.9	48.4	3.4	4.2	4.0	5.1	0.82	0.78
Management of companies and enterprises	6.1	3.0	1.3	1.5	1.2	1.3	0.85	0.89
Administration and support, waste management and remediation services	58.9	71.6	4.3	4.2	5.8	6.0	1.01	0.96
Ambulatory health care services	82.5	74.1	2.5	2.6	3.8	3.8	0.95	1.02
Other educational and health services	28.5	47.5	7.8	7.4	8.6	9.1	1.05	0.94
Arts, entertainment, and recreation	150.0	61.9	0.7	1.0	1.6	1.4	0.72	1.14
Food services and drinking places	13.8	35.3	7.4	6.0	7.2	6.7	1.24	1.07
Other accommodation and food services	22.0	11.1	0.8	1.5	0.8	1.4	0.54	0.61
Other services	18.5	27.5	4.8	3.9	4.9	4.1	1.24	1.19
Federal government	–28.9	–14.6	2.5	2.9	1.5	2.1	0.86	0.73
State government	6.1	15.8	3.5	3.9	3.2	3.8	0.89	0.84
Local government educational services	36.9	31.5	4.0	5.4	4.6	5.9	0.74	0.79
Other local government	24.6	22.6	3.3	4.6	3.5	4.7	0.71	0.74

SOURCE: BLS, using the NAICS industrial classification system and new 13-county Louisville MSA definition.

Economic Growth in Middle Tennessee: Can Local Public Services Keep Up?

David A. Penn

The Middle Tennessee economy is growing rapidly, creating opportunities for workers and businesses but also difficult challenges for local governments. With a population of 2.2 million in 41 counties, Middle Tennessee leads the state in population and job growth. Manufacturing is a very important source of jobs and payroll, with concentrations in transportation equipment, electrical equipment and appliances, and printing and publishing. Housing construction is expanding rapidly in response to population increases and income growth. Rapid growth, however, has strained local government services, challenging the ability of schools to accommodate growing numbers of students. Revenue options for local governments are limited, with most local revenues produced from property and sales taxes.

Federal Reserve Bank of St. Louis *Regional Economic Development*, 2005, *1*(1), pp. 30-39.

Middle Tennessee is growing rapidly: Two Midstate counties (Rutherford and Williamson) rank among the top 100 nationally in population growth, and Rutherford County is the fastest-growing large county in the United States for the third consecutive quarter as of 2004:Q4. As an informal definition, Middle Tennessee consists of 41 counties stretching from the Tennessee River in the west to the Cumberland Plateau in the east (Figure 1). The Midstate[1] population is 2,235,000 as of July 1, 2004; if it were a separate state, Middle Tennessee would rank 36th in size, smaller than Nevada but larger than New Mexico. The six largest counties account for more than half of the Midstate's population:

- Davidson County (572,000),
- Rutherford County (210,000),
- Williamson County (146,000),
- Montgomery County (142,000),
- Sumner County (141,000), and
- Wilson County (97,000).

The recently redefined Nashville-Davidson-Murfreesboro metropolitan statistical area (Nashville MSA) comprises 13 counties, with Davidson County at the center. Approximately two-thirds of the Midstate's population reside in the Nashville MSA. Population growth is relatively vigorous in the Nashville MSA, rising by 6.0 percent from 2000 to 2004, compared with 4.0 percent for the remaining Midstate counties, 2.4 percent for the rest of Tennessee, and 4.1 percent for the United States (Figure 2).

This study reviews the performance of the Middle Tennessee economy since the recent recession and examines the impact of rapid growth on the demand for public services and the consequent search for alternative revenue sources. The first section examines the industrial structure of the Midstate economy and summarizes recent trends in employment and housing construction.

[1] This paper uses the terms "Midstate" and "Middle Tennessee" interchangeably.

David A. Penn is the director of the Business and Economic Research Center and associate professor of economics and finance in the Jennings A. Jones College of Business at Middle Tennessee State University.

Figure 1

Counties in Midstate Tennessee

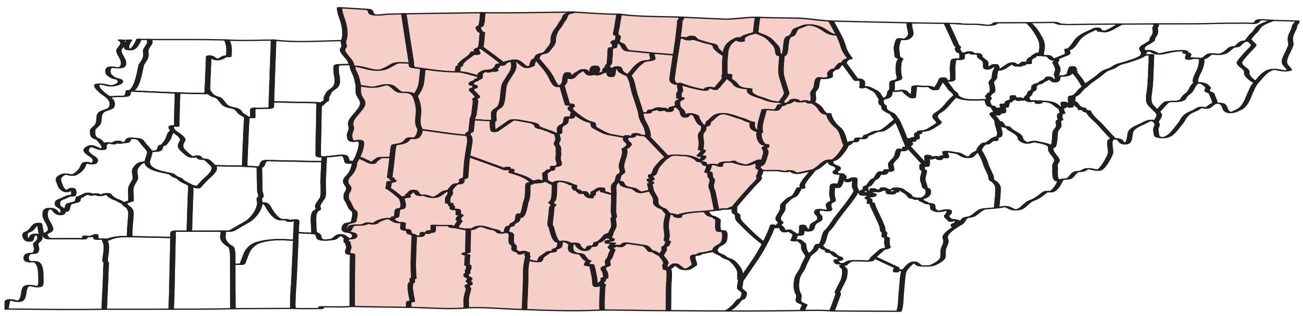
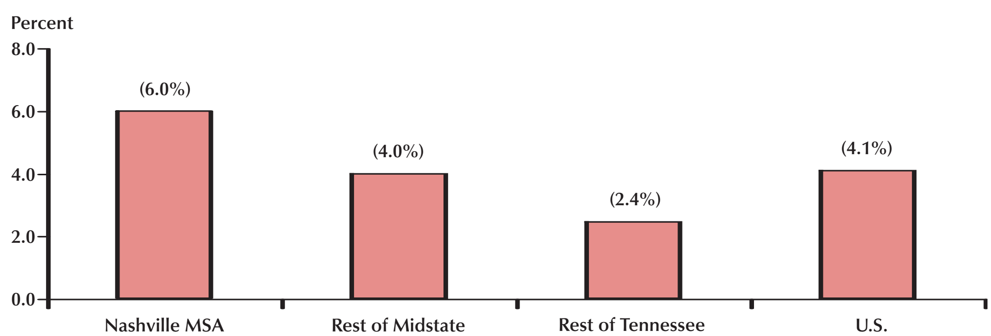


Figure 2

Population Growth 2000-04



The paper then sketches the effects of growth on the demand for public services, followed by a discussion of possible revenue sources for local governments. The final section offers conclusions.

INDUSTRIAL STRUCTURE AND RECENT ECONOMIC PERFORMANCE

Compared with the nation's economy, Tennessee is much more reliant on manufacturing as a source of payroll and employment. In fact, Tennessee has the ninth largest proportion of payroll from manufacturing among the 50 states. On average in the Midstate, manufacturing accounts for 16.5 percent of employment and 19.5 percent of total wages, much higher than the national average of 11.1 percent of employment and 13.6 percent of total wages.[2] But the Midstate average is heavily skewed by large Davidson County, which is much less dependent on manufacturing. Excluding Davidson County, the importance of manufacturing in the Midstate rises considerably to 22.8 percent of employment and 29.2 percent of total wages. In fact, manufacturing

[2] These figures are based on the QCEW (Quarterly Census of Employment and Wages) compiled by the Tennessee Department of Labor & Workforce Development for the U.S. Bureau of Labor Statistics.

generates at least one-third of total payroll for 17 of the 41 Midstate counties.

The comparative industrial structure of a local economy can be assessed using location quotients (LQs). An LQ is the share of employment (or wages) for a particular local industry divided by its national share. Thus, a Midstate industry with an LQ of more than 1 indicates that the sector is a more important source of jobs locally than nationally. Table 1 shows Midstate LQs for employment and total wages by major sector. With an LQ of 1.4, manufacturing clearly is much more important for the Midstate economy than for the nation. Two other major sectors, education and health services and leisure and hospitality, are also more important sources of wages in the Midstate than nationally. The high LQ for education and health services can be explained by a concentration of health care providers, headquarters for health care companies, and private universities in the Nashville MSA. The relatively large LQ for leisure and hospitality is due to country music, entertainment, and convention-related employers in Nashville.

Three manufacturing industries are particularly important in the Nashville MSA: transportation equipment manufacturing, electrical equipment and appliance manufacturing, and printing and related support activities; more than one in three manufacturing workers are employed in these industries. The transportation equipment sector includes large employers such as Nissan North America, Saturn, Bridgestone-Firestone, Visteon, and Peterbilt Motors. Numerous automotive parts manufacturers are also located in the Midstate. As for the printing industry, Nashville is well-known for its concentration of religious-oriented printing and publishing establishments.

Turning now to recent employment trends, the Midstate and national economies were affected similarly during the first 12 months of the 2001 recession. Employment for both peaked in the first quarter of 2001, reaching a trough about a year later. From peak to trough, payroll employment fell 1.8 percent for the national economy and 1.5 percent for the Midstate. By contrast, employment dropped much more for Tennessee, falling by 2.4 percent during the same interval.

Table 1

Midstate Location Quotients

Supersector	Employment	Total wages
Construction	0.408	0.500
Manufacturing	1.419	1.375
Trade, transportation, and utilities	1.010	1.043
Information	0.951	0.800
Financial activities	0.874	0.822
Professional and business services	0.929	0.814
Education and health services	1.007	1.194
Leisure and hospitality	1.004	1.132
Other services	0.804	0.885

NOTE: Calculated from QCEW series, third quarter 2004.

Since the end of the recession, however, Midstate job growth greatly outperformed that of the nation and state, rising 3.4 percent from first quarter 2002 to first quarter 2004 compared with 0.2 percent for the United States and 1.4 percent for Tennessee (Table 2). The much stronger job growth for the Midstate can be attributed to two factors: (i) employment in services-providing industries increased much more rapidly, particularly in the non-Nashville MSA counties, and (ii) job losses in the Midstate goods-producing sectors were less severe, thus exerting much less of a negative drag on net job growth. In fact, the Nashville MSA actually produced a modest job gain for the goods-producing sectors during the period.

As shown in Table 3, the strength of Midstate services-providing job growth from 2002 to 2004 is due to substantial gains in education and health services (9.6 percent), leisure and hospitality (6.6 percent), and trade, transportation, and utilities (6.3 percent). The information and manufacturing sectors lost jobs, while jobs were added in the construction sector.

Recent employment growth paths underscore the superior economic performance of the

Table 2

Employment Growth, 2002:Q1 to 2004:Q1

Area	Goods-producing	Services-providing	Private sector	Total
Midstate	–2.2%	5.6%	3.6%	3.4%
Nashville MSA	0.4%	4.7%	3.9%	3.8%
Tennessee	–3.7%	3.0%	1.3%	1.4%
U.S.	–4.7%	1.4%	0.1%	0.2%

NOTE: Calculated from QCEW series, U.S. Bureau of Labor Statistics.

Table 3

Midstate Employment Growth by Industry, 2002-04

Industry	Employment growth
Construction	3.1%
Manufacturing	–3.6%
Trade, transportation, and utilities	6.3%
Information	–8.6%
Financial activities	3.2%
Professional and business services	4.4%
Education and health services	9.6%
Leisure and hospitality	6.6%
Other services	1.8%

NOTE: Calculated from QCEW series, U.S. Bureau of Labor Statistics for the first quarters, private sector only.

Midstate. Figure 3 shows indexed payroll employment trends for the Midstate, Nashville MSA, Tennessee, and the United States. Beginning in the first quarter of 2003, employment rose more rapidly for the Midstate than the United States and the state overall.

Trends for manufacturing employment are similar, except that the Midstate and Tennessee itself peaked about one quarter before the nation did (Figure 4). Gains are strongest in the Nashville MSA; since the third quarter of 2003, employment rose 2.7 percent, a gain of 1,700 jobs, compared with 0.7 percent for the other Midstate counties, 0.6 percent for the rest of Tennessee, and –0.4 percent for the United States.

Employment growth in the services-providing industries is much stronger for the Midstate than for the state and nation (Figure 5). In the Midstate, services-providing sectors increased 5.6 percent from third quarter 2002 to first quarter 2004, compared with 3.0 percent for Tennessee and 1.4 percent for the United States.

Midstate Growth in Comparison

The Nashville MSA generated eight of every ten net new jobs in the Midstate since the first quarter of 2002 and six of ten net new jobs in Tennessee. As for individual counties, employment and wage growth vary considerably. Although some counties experienced very rapid growth during the past two years, others suffered losses. In fact, of the 41 counties comprising the Midstate, 13 counties experienced employment losses from third quarter 2002 to third quarter 2004. On the other hand, employment grew by a relatively strong 3.0 percent or more in 19 counties during the two-year period.

Each quarter the Bureau of Labor Statistics reports employment and wage growth for the 300 largest counties in the United States. The most recent report for the fourth quarter of 2004 shows Rutherford County ranking first nationally in terms of the employment growth rate, measured from December 2003 to December 2004. Rutherford County's payroll employment grew 8.9 percent over the year, outdistancing second-place Manatee County, Florida (8.7 percent), and third-place Clark County, Nevada (7.2 percent) (Bureau of Labor Statistics, July 19, 2005). In 2004, Rutherford County ranked first in three of four

Figure 3

Payroll Employment Trends for the Nashville MSA, Midstate, Tennessee, and the U.S.
(index of seasonally adjusted figures, 2000 = 100)

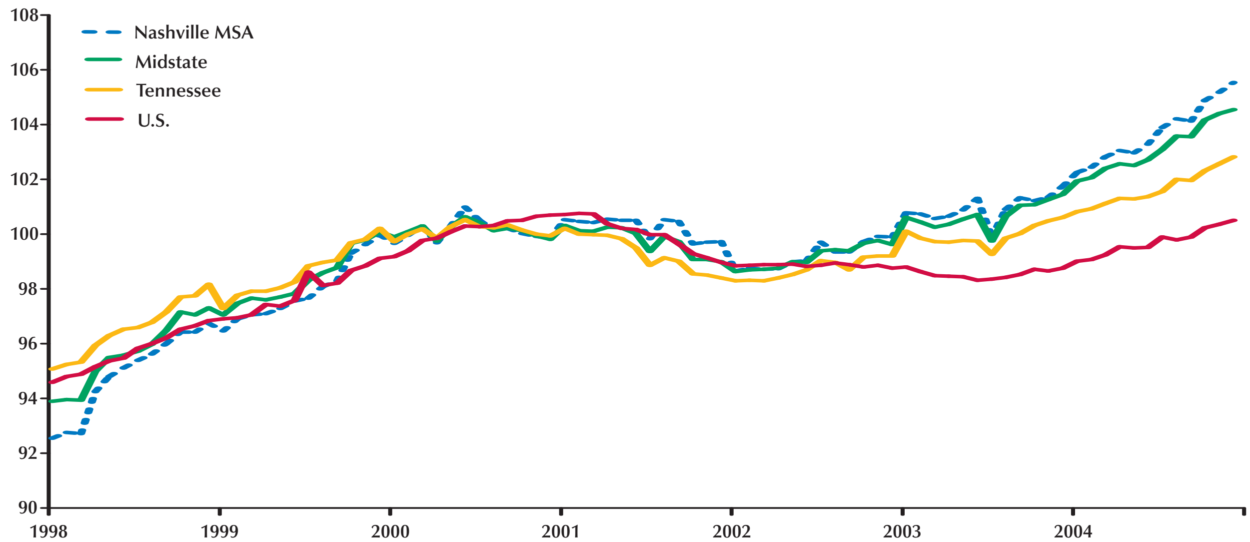


Figure 4

Manufacturing Employment Trends for the Nashville MSA, Midstate, Tennessee, and the U.S.
(index of seasonally adjusted figures, 2000 = 100)

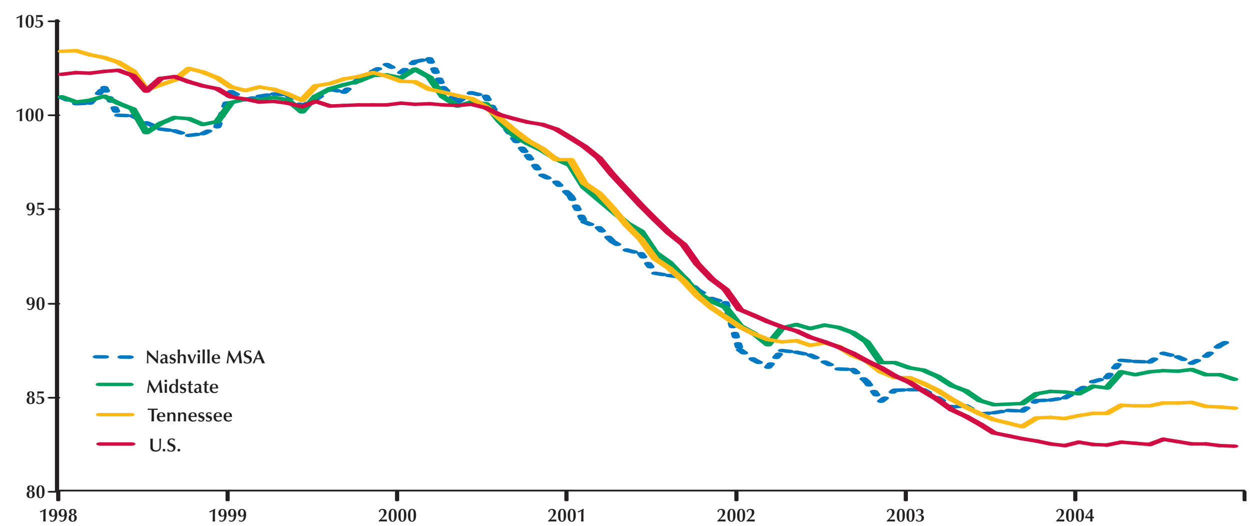
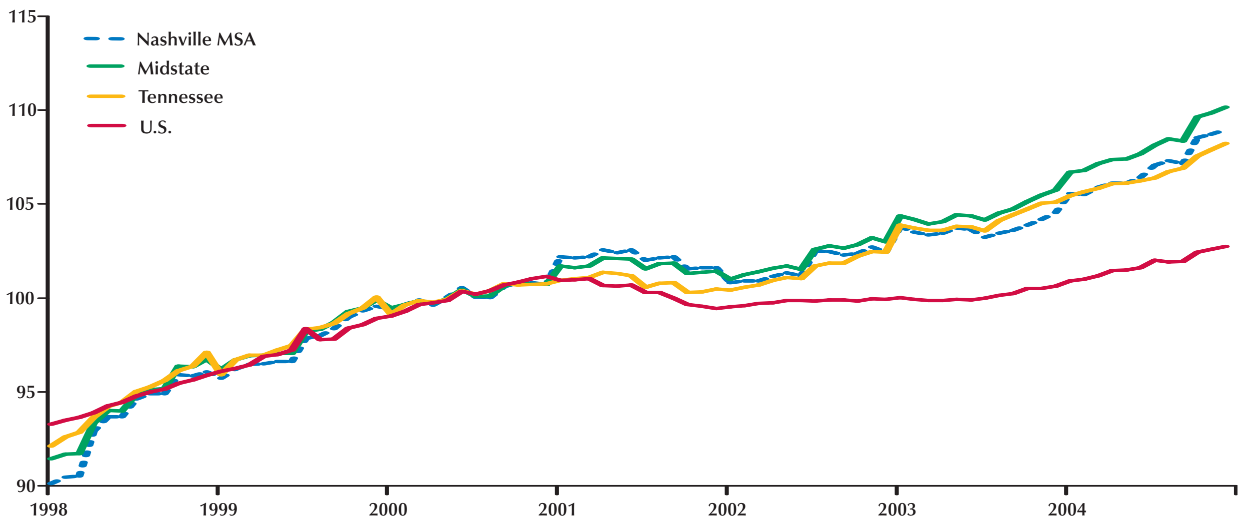


Figure 5

Services-Providing Industries Employment Trends for the Nashville MSA, Midstate, Tennessee, and the U.S.
(index of seasonally adjusted figures, 2000 = 100)



quarters in 2004 and ranked in the top ten growing counties for all of 2003 and 2004.

Examining the data more closely, we find that a very large share of Rutherford County's job growth, about 42 percent of growth during the past two years, occurred in the professional and business services sector. Virtually all the job growth in this sector appears to have originated from either temporary help agencies or companies that provide business support services. Similar job-growth patterns are evident among several of the other fastest-growing Midstate counties.

Very rapid employment growth in temporary help agencies and support services could be a sign that employers wish to hire but choose to proceed with caution; consequently, they hire temporary workers who can be easily laid off if business conditions suddenly turn sour. If businesses become convinced that growth is sustainable, they may hire more permanent workers and fewer temporary workers over the long run.

An alternative interpretation is that the temporary employment gains are not temporary but permanent. According to this interpretation, employers are out-sourcing certain needed skills and functions; growth of businesses that provide temporary employment services and support services for businesses could be tapping into a strong and growing demand for outsourced services, a demand that will likely continue as the national and worldwide economies become more and more competitive and pressures to improve productivity and minimize cost become even more intense.

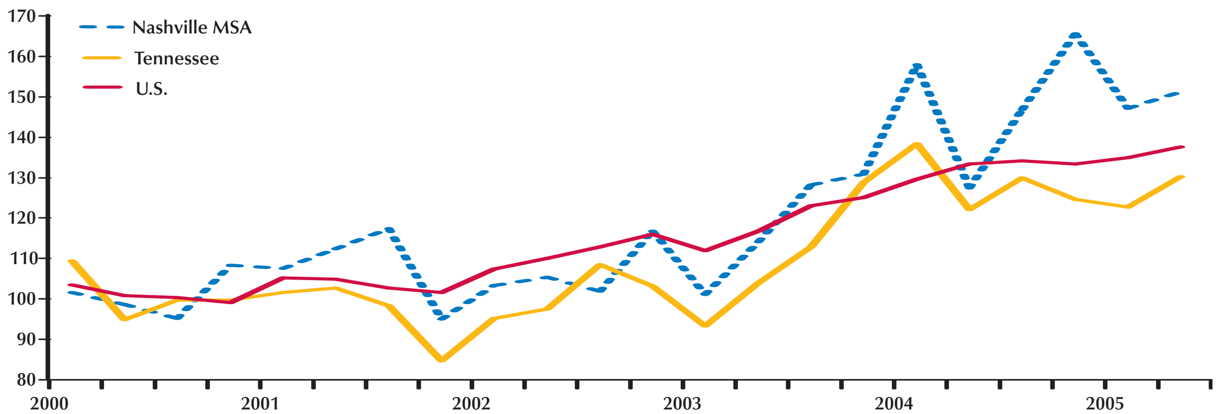
Housing Construction

New housing construction facilitates the movement of population into the Midstate area, particularly to the counties that form the outer ring of the Nashville MSA, including Rutherford, Wilson, Sumner, and Williamson. In the first quarter of 2005, Nashville MSA single-family home construction reached 3,300 units, a modest 4.0 percent gain over 2004.

Spurred by population growth, housing construction in the Nashville MSA is rising much faster than in the state or nation (Figure 6). As of the first quarter of 2005, Nashville MSA permit-

Figure 6

Index of Single-Family Home Construction Permits, 2000:Q1 to 2005:Q2
(index of seasonally adjusted figures, 2000 = 100)



authorized single-family home construction reached 149 percent of its level from the first quarter of 2000, compared with 129 percent for Tennessee and 128 percent for the nation.

Housing construction creates well-paying jobs for people in a number of specialized occupations including carpenters, floor layers, roofers, plumbers, electricians, and so on. Construction also creates additional demand for building materials such as brick, concrete, stone, lumber, and steel; many building materials are purchased locally by construction contractors. From third quarter 2000 to third quarter 2004, employment in the residential building construction sector (North American Industry Classification System [NAICS] 2361) increased 31.2 percent in the Nashville MSA, rising from 3,000 in 2000 to 3,900 in 2004. Total wages paid in this industry jumped by 47.8 percent during the same period. By 2004, the average employee in the industry earned $40,100 during the first quarter on an annual basis, far higher than the average Nashville MSA across-industry pay.

Relatively recent evidence suggests that slower growth may be in the cards for housing construction in the Nashville MSA. Interestingly, the problem is not demand, because demand is expected to remain strong as long as mortgage rates remain at relatively low levels. Rather, construction activity may slow because of supply-side bottlenecks caused by an acute shortage of lots to build on. Land developers are hard pressed to keep up with the demand for lots by the construction industry. Tighter supply coupled with unabated demand has resulted in higher prices for developed lots, on the order of 10 percent to 20 percent over the year. One builder reports that higher costs for lots have added $4,000 to $10,000 to the price of homes he builds.

Greater scrutiny from local land-use planners and related environmental concerns have added several months to the process of getting a new development approved; a year or more is now required for approval, compared with six months in the recent past. Also, developers and builders are moving from easier-to-develop properties to areas that are more difficult to develop. Because the prime (and lower-cost) properties have already been developed, builders are moving further out from the central locations into areas that are more challenging because of rock content, lack of infrastructure, and other engineering difficulties (Russell, 2005).

Eventually, higher overall home prices due to higher development costs and supply bottlenecks may slow the rate of growth of housing construction. Little evidence exists to show this is occurring presently, however, as closings on homes in the Nashville MSA remain strong, median prices are rising, and the inventory of unsold homes is falling.

DEMAND FOR PUBLIC SERVICES

Rapid employment and population growth produce greater economic opportunities for both businesses and workers. For example, the expansion of the local employment base means that Midstate high school and college graduates have a much better chance of pursuing a career close to home. However, growth can also strain the ability of local governments to provide necessary services. Local governments must provide critical services such as police and fire protection, K-12 education, streets and roads, water supply, sanitary sewers, and waste disposal for a growing population. In general, rapid growth not only raises the demand for local government services, but may also increase the average cost of providing the services and affect local quality of life. For example:

- Wastewater treatment: Costs are rising because the local watershed is unable to accept additional treated wastewater. Some local governments are beginning to rely on nontraditional, and more expensive, methods of disposal such as land spraying.
- Air quality: Increased consumption of gasoline related to traffic growth exacerbates local air-quality problems. Reduced air quality can cause health problems in at-risk populations and lost productivity for workers. State and local officials are working to develop alternatives to slow the rate of growth of vehicle emissions.
- Higher education: Two local public universities are experiencing rapid enrollment growth, straining the institutions' ability to maintain quality.
- Law enforcement: Some jails are overcrowded, and some sheriffs complain about lack of funding. Some have gone so far as to sue county governments for additional funds.

Public Schools

The effect of growth on the demand for K-12 education services in the Midstate is dramatic. School systems in high-growth areas must accommodate increasing numbers of families with children moving into the area. Maintaining class sizes at the present level means that schools must hire, on average, one additional teacher for every 15 additional K-12 students. Further, additional support staff, counselors, and administrators are needed to maintain and manage a new or expanded school.

On the cost side, operating costs per student in the Midstate rose from $5,500 to $6,700 per student over the most recent four years, a 5.2 percent annual rate of growth, considerably greater than the rate of inflation as measured by the growth of the consumer price index. Even school systems with slow enrollment growth are experiencing substantial cost increases. For example, enrollment in Nashville city schools rose just 1.4 percent from 2000 to 2004, yet operations expenditures increased 30.4 percent (Tennessee Department of Education).

The Nashville city government is considering a proposal to increase the sales tax rate from 9.25 percent to the state maximum 9.75 percent, with most of the revenue intended for rising costs of providing education. Increased costs for Nashville schools are driven by four factors: seniority raises required by contract, pension and insurance benefit costs, expenses related to opening new schools, and annual pay increases (Long, 2005).

In the fall of 2000, public schools in the Nashville MSA enrolled about one fifth of Tennessee's schoolchildren and Midstate schools enrolled approximately 36 percent. During the 2000-04 period, the vast majority of Tennessee's net enrollment growth occurred in the Midstate counties: Average daily attendance in Tennessee public schools rose by a total of 1.7 percent, but in the Midstate counties it rose by 4.1 percent and in the Nashville MSA by 6.1 percent. Put another way, of the net new 14,314 children in Tennessee public schools, 77 percent were in the Nashville MSA and 89 percent in the Midstate counties.

Rutherford County experienced the highest enrollment growth in the state, both in absolute terms and the growth rate: The county school system added 4,100 students during 2000-04, rising by 4.2 percent annually. Other Midstate counties with enrollment growth of 1,000 or more include Williamson, Wilson, Sumner, and Montgomery. Taken together, these five counties experienced an aggregate enrollment increase of 11,000 students, about three quarters of Tennessee's net increase during the four-year period.

The demand for classroom space, teachers, and all the materials and supplies needed to operate schools caused spending for operations to increase greatly: From 2000 to 2004, Midstate schools increased spending from $1.728 billion to $2.175 billion, a 25.8 percent increase. Interestingly, the revenue stream required to pay for this increase relied mostly on local sources. Total revenue received from the state increased just 11.5 percent from 2000 to 2004. By contrast, total local revenue rose by 35.8 percent, mostly from increased property tax collections due to housing and commercial growth and also to higher property tax rates.

Capital expenditures are also on the rise. The Rutherford County school system, for example, estimates that, if present enrollment trends continue, the county will need to build two to three new schools each year for the next decade, an estimated expenditure of more than $500 million plus millions more for annual operating expenses to hire new teachers and staff.

SOURCES OF REVENUE FOR LOCAL GOVERNMENTS

The rapidly rising demand for public services, especially schools, has produced a lively discussion of how to pay for growth. Options for local government, especially county government, are not numerous (Penn, 2004). County governments rely on three primary sources: state revenue, local property taxes, and local option sales taxes. Other revenue sources exist, such as development fees, wheel taxes, and adequate facilities taxes, but these generate much smaller revenue streams. Development fees and adequate facilities taxes recover at least part of the additional costs associated with providing public services to new residents. A development fee is a flat fee charged on each new housing unit; the development fee is currently $1,500 in Rutherford County. By contrast, the adequate facilities tax is a charge levied per square foot of a new home ($1.00 per square foot, for example). Thus, the tax levy varies from home to home depending on square footage—larger homes pay more, smaller homes pay less. Of the two, the development fee is more regressive; a high-priced new home pays a smaller percentage of the sales price to the development fee than does a low-priced new home.

The wheel tax is an annual fee collected for each vehicle owned. The wheel tax ranges from $25 to $50 per vehicle in Midstate counties. In Rutherford County, a $40 wheel tax generated $6.6 million in revenue in fiscal year 2002, making the tax the third-largest source of local revenue for the county government.

Not surprisingly, local residents are reluctant to raise taxes; a recent exception is Dickson County, where voters approved a $20 increase in the wheel tax in January 2005. By contrast, voters in Williamson County recently defeated a tax-increase proposal.

The local option sales tax is a very important source of funds for cities, schools, and county government. Some counties, such as Rutherford County, currently levy the maximum sales tax rate allowed by state law, a combined state and local rate of 9.75 percent on most items. Interestingly, the spending base for the state sales tax and the base for the local option sales tax are not the same. The most important difference is that spending subject to the local option tax has a single-article limit; the local portion of the sales tax applies only to the first $1,600 of the sales price for a single article. For example, the entire value of a $2,000 plasma television is subject to the 7 percent state sales tax rate, but the local option sales tax is limited to the first $1,600 of the transaction. When the article is a $35,000 vehicle, it is easy to see that the state treasury collects much more revenue per penny of tax than does the local government. When the economy is growing and big-ticket items such as vehicles, furniture, and large appliances are selling well, local governments receive less of a revenue boost from the tax on sales, due to the single-article cap, than does the state government.

Perhaps the biggest hurdle facing local governments, particularly county governments, in fast-growing areas is how to expand the portfolio of revenue options available to pay for rising costs of services and growth-related costs and at the same time avoid property tax increases. Explicit approval from the state legislature is necessary to implement new kinds of taxes or to increase certain fees or taxes. For example, a county cannot on its own authority levy a development tax or adequate facilities tax without the specific permission of the state legislature; increases for some existing taxes and fees typically must also pass muster with the legislature. This requirement creates obvious opportunities for lobbying efforts from opponents of growth-related taxes. For example, a local real estate transaction tax proposed in the legislature several years ago was defeated in the legislature after intensive lobbying efforts by the real estate industry.

The housing construction and real estate industries argue that increasing the development tax or levying new adequate facilities taxes places too much of the burden on a relatively narrow portion of the housing market—new housing. The Rutherford County home-building industry has stated that it would consider supporting a broad-based tax, such as a local real estate transaction tax, since it would apply to sales of both new and existing homes. The real estate brokerage industry, however, remains opposed to any new taxes (Shaw, 2004).

Some communities have considered placing limits on growth. For example, city leaders in Franklin (Williamson County) recently considered a temporary moratorium on any new zoning changes that would allow additional housing growth. Interestingly, the Franklin city administrator warned that limiting the annual growth of housing below 600 single-family units could have very significant impacts on the city's budget. The city has substantial debt service obligations related to new wastewater treatment capacity; if fees and tax revenue from new housing construction are not sufficient to cover the annual debt service, other city services must be cut to make up the difference (Watson, 2005).

CONCLUSIONS

The Midstate economy generates a large number of jobs in a variety of industries. Rapid job and population growth create opportunities for workers and businesses, but also produce considerable stress on the ability of local governments to provide services. Local revenue sources have difficulty keeping pace with growth if there are no tax increases. Counties rely on the property tax to generate funds that are not sufficiently forthcoming from other sources such as the local option sales tax. Voters are voicing more and more concern about steadily rising property taxes, forcing local governments to more aggressively pursue alternative sources of revenue.

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Income Inequality in Rural Southeast Missouri

Bruce Domazlicky

Income inequality has been increasing in the United States since at least 1980. However, in a 34-county region of southeast Missouri, income inequality actually decreased from 1990 to 2000. As well, income inequality was less in the selected region as compared with the entire United States in 1999. A simple, single-equation regression model is used to identify the factors that influence income inequality in southeast Missouri. Five factors stand out as especially significant: the percent of the population under 18 years of age, the percent of the families that are female-headed, the female labor force participation rate, the level of income, and the percent of the population with a high school diploma (but no higher degree). Income inequality increases with income and the percent of female-headed families, whereas it decreases with increases in the other three factors.

Federal Reserve Bank of St. Louis *Regional Economic Development*, 2005, *1*(1), pp. 40-51.

There is strong evidence that income inequality in the United States has been increasing since at least 1980 (see, for example, Levernier, 1996; Levernier, Partridge, and Rickman, 1995 and 1998a; Partridge, Rickman, and Levernier, 1996). In addition, there is considerable variation in income inequality at the regional level. For example, income inequality tends to be higher in non-metropolitan areas than it is in metropolitan areas (Levernier, Partridge, and Rickman, 1998b). The causes of the observed regional variation have been studied by researchers for states (Levernier, Partridge, and Rickman, 1995; Partridge, Rickman, and Levernier, 1996), for counties (Levernier, Partridge, and Rickman, 1998a), and for urban areas (Garafalo and Fogarty, 1979).

Although some studies have considered the entire population of over 3,000 counties in the United States (for example, Levernier, Partridge, and Rickman, 1998a), this study considers income inequality in a small sample of 34 counties in southeast Missouri at two points in time: 1989 and 1999. We are concerned with identifying the causes for the variation in income inequality that exists in the 34 counties and for the changes that occur over the 10-year period. Studying a subset of the population of 3,000 U.S. counties allows us to determine whether the factors affecting income inequality in our small sample are similar to those in the entire population of counties. When the 3,000 counties are included in a single study, some of the unique characteristics of the many subregions in the United States are surely lost because of simple aggregation. Therefore, there is value in bringing the microscope to bear on a small region of the entire country. In addition, because income inequality changes over time, it is important to identify those factors that have a continuing effect on inequality as opposed to factors that have a more transitory effect.

Although the measurement of income inequality and the identification of the factors that influence inequality are interesting endeavors in their own right, the ultimate goal of a study such as this one must be to make policy prescriptions

Bruce Domazlicky is a professor of economics at Southeast Missouri State University.

that emanate from the results. Persistent and increasing income inequality is likely to be deemed undesirable by many in society, though that can be a controversial statement. In fact, Kuznets (1955) was one of the first to point out that income inequality may increase initially as a region or country develops. That is, it may be normal or even necessary over a period of time for income inequality to increase as income progresses. This implies, perhaps, that attempts to reduce income inequality could be futile or even harmful. Nevertheless, economic policies that reduce inequality are likely to be favored over those that increase it. The results from this study are likely to be useful in crafting policies that promote greater income equality in southeast Missouri.

The organization of this paper is as follows. A review of income inequality in rural counties in southeast Missouri is given in the next section. The third section introduces the basic model that is used in this paper to identify the factors that affect the variation in income inequality in rural southeast Missouri. In this section, we also briefly review results from selected earlier studies. The fourth section outlines the results of the model for southeast Missouri. The final section offers a brief summary and conclusion.

INCOME INEQUALITY IN SOUTHEAST MISSOURI

The Gini coefficient, a simple measure of income inequality with a value that ranges from 0 (no inequality) to 1 (complete inequality), was used in this study as the measure of income inequality in a county. The U.S. Census Bureau (historical income inequality tables; www.census.gov/hhes/www/income/histinc/f04.html) reports that the Gini coefficient for the entire nation was equal to 0.401 in 1989 and 0.429 in 1999—evidence of rising income inequality in the United States.

Thirty-four rural counties in southeast Missouri comprise the sample. (See Figure 1 for a county map of the state of Missouri.) The 1990 U.S. Census (summary file 3; www.census.gov/main/www/cen2000.html) reports the total family income for a county as well as the number of families in 25 different income classes. The 2000 U.S. Census reports total income and number of families for 16 different income classes. Using the method of previous researchers, we assume that income level for each family is equal to the midpoint of its income class. For the highest, open-ended income class, we assume families are at the mean of the income class as reported by the U.S. Census Bureau.

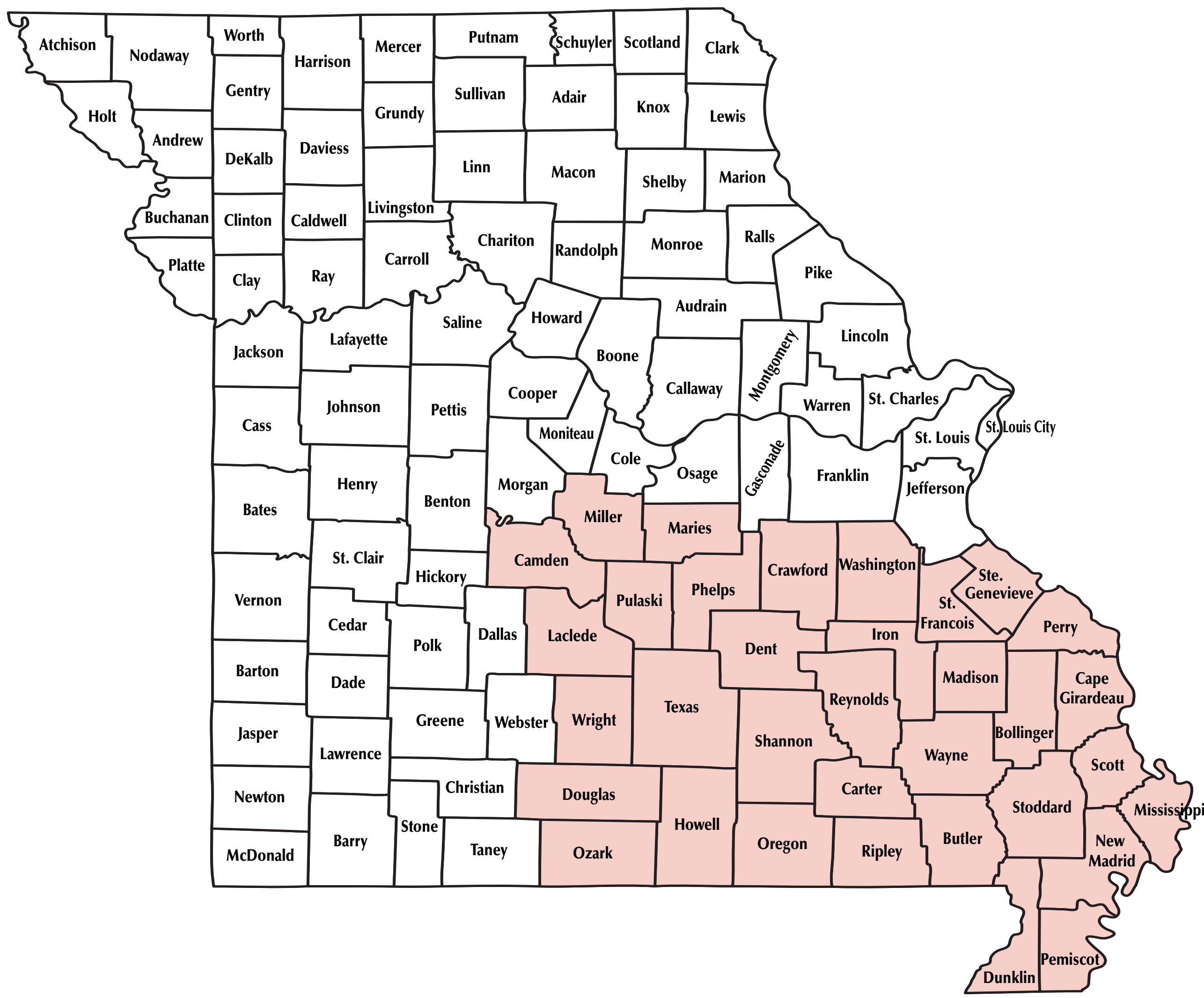
Maxwell (1990) explains the actual procedure for estimating the Gini coefficient based on income class data. We followed this procedure to arrive at the results in Table 1, which gives the Gini coefficient for each county in our study for 1989 and 1999. Note that the income collected in a decennial census is actually for the previous year; therefore, the Gini coefficients technically are for 1989 and 1999. Comparison data for the United States and for all of southeast Missouri are given in Table 1 as well. Note that the southeast Missouri data are not computed as the average of the 34 counties, but, rather, they are computed using the aggregate data for the entire region.

Two facts are readily apparent from the data in Table 1. First, income inequality in southeast Missouri is less than that in the entire United States in 1999. Second, income inequality actually fell in southeast Missouri in the 1990s, while it was rising in the United States. Income inequality fell in 19 of the 34 counties, rose in 14 counties, and remained unchanged in 1 county (Cape Girardeau). The Gini coefficient ranged from 0.3421 to 0.4815 in 1989. In 1999, the coefficient ranged from 0.3366 to 0.4809. One reason for the lower level of income inequality in southeast Missouri could be the lower overall rate of growth in the region relative to the rest of the country and relative to urbanized regions.

Although there are significant changes in levels and rankings from 1989 to 1999, the simple correlation coefficient between the two years is 0.69, which is highly significant at the 1 percent level. Counties in the Bootheel region of Missouri (such as Pemiscot, Dunklin, Mississippi, and New Madrid) have among the highest Gini coefficients. Counties just north of the Bootheel (Cape Girardeau, Ste. Genevieve, Perry, St. Francois,

Figure 1

Missouri Counties



and Bollinger) generally have lower Gini coefficients. Counties in the Ozarks region are more of a mixed bag, with some below the average (e.g., Laclede, Maries, and Miller) and others above the average (e.g., Wayne, Butler, Ripley, Douglas, and Howell). Therefore, it will be important in our empirical analysis to adjust for potential regional effects that might exist.

Table 1

Gini Coefficients

County	Gini 1989	Rank 1989	Gini 1999	Rank 1999
United States	0.4010		0.4290	
Southeast Missouri	0.4128		0.4097	
Bollinger	0.3816	28	0.3586	30
Butler	0.4306	8	0.4477	2
Camden	0.4097	13	0.4152	15
Cape Girardeau	0.3834	26	0.3834	26
Carter	0.3993	16	0.4207	13
Crawford	0.3819	27	0.3883	25
Dent	0.4360	7	0.3935	23
Douglas	0.4476	4	0.3956	21
Dunklin	0.4618	2	0.4262	10
Howell	0.4264	9	0.4127	16
Iron	0.3970	19	0.4303	5
Laclede	0.3695	31	0.3962	20
Madison	0.3808	30	0.3992	19
Maries	0.3897	22	0.3496	32
Miller	0.3809	29	0.3673	29
Mississippi	0.4458	5	0.4450	3
New Madrid	0.4424	6	0.4292	6
Oregon	0.4238	10	0.4256	11
Ozark	0.3889	23	0.4281	8
Pemiscot	0.4481	3	0.4809	1
Perry	0.3532	32	0.3455	33
Phelps	0.4064	14	0.3942	24
Pulaski	0.3421	34	0.3360	34
Reynolds	0.3954	20	0.3812	28
Ripley	0.4199	12	0.4170	14
St. Francois	0.3853	25	0.3816	27
Ste. Genevieve	0.3464	33	0.3524	31
Scott	0.3989	17	0.3951	22
Shannon	0.3949	21	0.4416	4
Stoddard	0.4224	11	0.4062	18
Texas	0.3977	18	0.4270	9
Washington	0.4002	15	0.4225	12
Wayne	0.4815	1	0.4290	7
Wright	0.3864	24	0.4064	17

NOTE: Spearman's rank correlation coefficient = 0.643 (t = 3.69).

MODEL AND LITERATURE REVIEW

The general approach to identifying the factors associated with income inequality is a simple model of the form:

$$INEQ = \mathrm{F}(DEMOG, LF, INDCOMP, HUMANK, GEOG),$$

where *INEQ* is some measure of income inequality in a region such as the Gini coefficient, *DEMOG* includes demographic variables, *LF* denotes variables related to conditions in the labor force, *GEOG* are variables that relate to regional effects, *INDCOMP* includes variables that measure the industrial composition of a region, and *HUMANK* are human capital variables. The various studies differ as to the exact variables that are included in each category and to the different categories that might be used. However, in all cases, economic theory is used to identify and to support the use of the individual variables that are included in the model.

The four demographic variables most often used are the percents of the dependent population that are under 18 (*UNDER18*) or over 64 years of age (*OVER64*) (two separate variables), the percent of African-Americans and/or other minorities (*BLACK*), and the percent of families that are headed by a female (*FEMALE*). Because of possible discrimination in the labor force, a greater proportion of African-Americans and/or other minorities in a region may lead to greater income inequality. This has generally been found to be true in previous studies (Persky and Tam, 1994; Levernier, 1996 and 1999; Partridge, Rickman, and Levernier, 1996). It is also expected that the greater the percent of the population that is dependent, the greater will be the degree of income inequality. People 65 years of age or older frequently have lower incomes. A greater proportion of this age group is likely to increase income inequality in a region. Similarly, the population under 18 usually receives little or no income, which could also contribute to income inequality. However, the actual research is mixed with respect to these variables. In some cases, just one group is found to be significant or has an unexpected sign.[1] Female-headed families are much more likely to be low income than are other families; therefore, as the percent of such families increases in a region, income inequality should increase. Most research finds this to be the case (see, for example, Levernier, Rickman, and Partridge, 1995 and 1998a).

Four types of variables fall into the labor force category. One variable relates to the labor force participation rate; here it is exclusively women (*FLFPR*). Women increased their participation in the labor force in record numbers starting in the 1970s, a trend that has continued through the 1990s. The entrance of women into the labor force will boost the earnings of the affected families and will contribute to reductions in income inequality if the women are from lower and middle class families. If women from upper middle income and upper income families enter the labor force, it is possible that increased labor force participation by women will increase income inequality. The overwhelming majority of studies find that income inequality falls when the labor force participation rate of women increases (Levernier, 1999; Levernier, Partridge, and Rickman, 1995 and 1998a). Instead of using the labor force participation rate for women only, some studies use the employment rate (Levernier, 1996) or the labor force participation rate for both sexes (Partridge, Rickman, and Levernier, 1996) with similar results. Our study includes only the female labor force participation rate.

A second labor force variable used in many studies is the percent of the population that is foreign born. Several studies find a positive and significant relationship between foreign born workers and income inequality (see, for example, Levernier, 1996). The theory is that foreign-born individuals frequently have lower skills or language impediments that reduce their income, thus contributing to income inequality. Because there are so few foreign born workers in the counties in our sample, this variable was not significant in any of the regressions and, therefore, is not included in our final regressions.

[1] Levernier (1999), for example, found only the group under 18 to be positively related to income inequality; Levernier, Partridge, and Rickman (1998a) found that as the percent of the population over 64 increases, income inequality decreases in nonmetropolitan counties.

A third labor force variable relates to the conditions of the labor market in a region. Increases in employment (*EMPGROW*) in a region offer opportunities for unemployed individuals to increase their incomes, which should help to lower income inequality (see, for example, Levernier, Partridge, and Rickman, 1995; Levernier, 1999). Therefore, employment growth in the previous decade is included in the model as a measure of employment opportunities in the region.

The final labor force variable is the income of the region. The Kuznets (1955) hypothesis indicates that income inequality may increase as income in a region increases initially and then may decrease as income increases further. Therefore, a region's level of income inequality may be influenced by its present stage of economic development. Levernier, Partridge, and Rickman (1998a), for example, find a positive relationship between income level and income inequality for their sample of over 3,000 counties. Bishop, Formby, and Thistle (1992) also find a positive relationship for income. They use states in 1980 for their sample. However, Persky and Tam (1994) find a negative relationship between income and income inequality. So, because the relationship between income inequality and the level of income may not be linear, two models were tested here in addition to a simple linear model. One was a quadratic approach on the level of income. The empirical results did not support a quadratic approach. The second approach was to use the log form for income (*LINCOME*). This proved more satisfactory and was adopted for the final model.

Industrial composition variables relate to the type of industries that are found in a region. One hypothesis is that a large manufacturing sector offers relatively high-wage employment to less-educated workers, thereby contributing to a reduction in income inequality. Conversely, if employment in a region is concentrated in the retail and/or service sectors, this could lead to increases in income inequality. Another sector that could be of importance in determining income inequality is farm employment. Farm income is notoriously variable and frequently low; both of these facts could lead to greater income inequality in regions with a large farm sector. The ideal approach, perhaps, is that used by Levernier (1999) or Levernier, Partridge, and Rickman (1998a). They include the percent of employment in each major SIC (standard industrial classification) sector. However, because of restricted degrees of freedom in a small sample, we tested only two variables: the percent employed in the manufacturing sector (*MFG*) and the percent employed in the farm sector (*FARM*). We also tried, as an alternative measure, the percent of regional income for these two industries, but neither was significant.

Three human capital variables relating to education of the labor force have been used in various studies. Two variables relate to the level of education: the percent of the population (25 years of age or older) that has a bachelor's degree or higher (*COLLEGE*) and the percent of the population that has a high school diploma (but no college degree) (*HS*). The latter category includes individuals with some college and/or an associate's degree. Therefore, the excluded category is high school dropouts. It is difficult to say, a priori, how more college graduates in a region may affect income inequality. It is possible that more college graduates will increase income inequality. An increase in the population with high school diplomas is likely to decrease income inequality. Levernier, Partridge, and Rickman (1995), for example, find that increases in the percent of college graduates increase income inequality, whereas increases in the percent of those with a high school diploma decrease it. In addition to the level of education, several studies use the standard deviation of educational attainment (*EDUC*) in a region. The U.S.Census Bureau reports the number of individuals in a region in each education category: less than eighth grade education, high school dropout, high school diploma, etc. We take the standard deviation of these reported groups. It is generally found that a greater dispersion of educational attainment increases income inequality (Levernier, Partridge, and Rickman, 1998a).

In addition to the variables that have been discussed thus far, it is also likely that other factors that influence income inequality are unique to given regions. In addition, there may be omitted factors that are not measured by the variables

Table 2

Variable Statistics

Variable	Mean	Standard deviation	Maximum	Minimum
GINI (x100)	40.42	3.22	48.1	33.6
OVER64	16.21	2.44	21.1	6.7
UNDER18	25.98	1.95	30.9	20.2
BLACK	3.38	6.37	27.3	0.0
FEMALE	7.95	2.86	18.6	3.2
FLFPR	48.87	5.43	61.5	37.8
EMPGROW	16.62	14.12	52.8	–24.6
LINCOME	10.00	0.14	10.3	9.7
LPOP	9.83	0.61	11.1	8.6
MFG	16.64	7.98	34.1	1.7
FARM	11.83	7.08	31.2	2.5
COLLEGE	9.51	3.69	24.2	5.8
HS	54.79	6.62	66.3	42.1
EDUC	12.60	1.31	15.8	8.9

included within the model. It is important to control for these regional effects and omitted factors, usually through the use of dummy variables. In our model, we have five dummy variables that relate to regional effects. The state of Missouri is divided into regional planning areas, each served by a regional planning commission. Our 34 counties fall into six different regional planning areas. We use five dummy variables for the following planning commission areas: the Bootheel, Lake of the Ozarks, Meramec, Ozark Foothills, and South Central Ozarks. The excluded area is the Southeast Regional Planning area, which includes seven counties. (See Appendix A.) Although it might be preferable to include a dummy variable for each of the 34 counties individually (minus one to avoid perfect collinearity), limited degrees of freedom do not favor such an approach.[2] Counties in planning areas are likely to be fairly homogeneous, rendering a planning area approach tenable. One additional variable relating to geography is the population (*LPOP*) of the county. Income inequality may be affected by economies of scale or agglomeration economies, which can be approximated by the population of the county. Levernier, Partridge, and Rickman (1998a) found that the log of population was negatively related to income inequality in metropolitan counties, but it was insignificant in nonmetropolitan counties. Because our sample includes solely rural counties, it is possible that mere population size may not have any discernible effect on income inequality.

A final variable to be included in our model is a dummy variable representing time (TIME). The variable is equal to 1 for 1999 and 0 for 1989. This variable will capture any unique time-specific characteristics for the two time periods that are not captured by other regressors in the models.

EMPIRICAL RESULTS

Variable definitions are given in Appendix B, and variable statistics are given in Table 2. Ordinary least-squares regression was used with

[2] We did try a model that included dummy variables for each of the counties. The results were virtually the same, except that the high school variable was insignificant. As noted later, the results for the high school variable exhibit considerable instability and are to be interpreted with care. Only three of the county dummy variables were significant at the 10 percent level, and the adjusted R^2 was only marginally higher. We decided to report the model with the planning commission dummies because it allowed for greater degrees of freedom.

Table 3

Empirical Results

Variable	(1)	(2)	(3)
CONSTANT	*53.886 (1.75)	54.24 (1.54)	–16.985 (0.47)
OVER64	0.150 (0.87)	0.150 (0.86)	0.053 (0.33)
UNDER18	**–0.508 (2.21)	**–0.510 (2.09)	***–0.671 (3.03)
BLACK	–0.014 (0.15)	–0.014 (0.15)	–0.047 (0.52)
FEMALE	***0.874 (3.63)	***0.877 (3.20)	***0.948 (3.87)
FLFPR	***–0.378 (4.11)	***–0.377 (3.91)	***–0.298 (3.44)
EMPGROW	0.014 (0.77)	0.014 (0.76)	–0.003 (0.15)
LINCOME	–0.888 (0.29)	–0.925 (0.26)	**8.262 (2.15)
LPOP	**1.515 (2.56)	**1.511 (2.44)	0.584 (0.91)
MFG	0.060 (1.59)	0.060 (1.57)	0.027 (0.71)
FARM	**0.137 (2.32)	**0.137 (2.29)	0.039 (0.63)
HS	–0.019 (0.24)	–0.019 (0.21)	**–0.191 (2.01)
COLLEGE	0.066 (0.51)	0.066 (0.51)	–0.119 (0.94)
EDUC	0.03 (0.09)	0.033 (0.09)	0.044 (0.13)
TIME		–0.024 (0.02)	1.975 (1.62)
BOOTHEEL			1.301 (1.33)
LAKEOZ			***2.459 (2.83)
MERAMEC			**1.970 (2.38)
OZFOOT			**1.879 (2.12)
SCOZ			***4.612 (4.36)
R^2 (adjusted)	0.69	0.69	0.76
F-statistic	12.70	11.58	12.18

NOTE: Dependent variable: *GINI*; estimation: least-squares regression; number of observations: 68; numbers in parentheses are absolute values of *t*-tests; */**/*** indicate statistical significance at the 10/5/1 percent levels, respectively.

the Gini coefficient (multiplied by 100) as the dependent variable. Three regressions are reported in Table 3: Regression (1) excludes the time dummy and the regional dummy variables, regression (2) adds the time dummy, and regression (3) adds the regional dummy variables. The inclusion of the dummy variable for time has no effect on the regression. The variable *TIME* is negative and not significant in regression (2) but changes sign and approaches significance at the 10 percent level in regression (3).

The inclusion of the regional dummy variables does have a significant effect on the regression as two variables lose significance (*FARM*, *LPOP*) and two become significant (*LINCOME*, *HS*). Three variables are highly significant in all three regressions (*FEMALE*, *FLFPR*, *UNDER18*). The partial F for the inclusion of the regional dummy variable is 4.25, which is significant at the 1 percent level. This means the regional dummy variables should be included in the model. Therefore, our remarks will pertain mainly to regression (3) in Table 3.

All of the coefficients on the regional dummy variables are positive and four are significant at the 5 percent level. Apparently, income inequality increases as we move away from the seven counties served by the Southeast Missouri Regional Planning Commission. Beyond the regional dummy variables, five independent variables are significant at the 5 percent level or better: *FEMALE*, *LINCOME*, *FLFPR*, *UNDER18*, and *HS*. No other independent variable is significant at even the 20 percent level.

Similar to most other studies, this study shows a positive and highly significant relationship (better than 1 percent level) between income inequality and the percent of families that are headed by a female. The low level of income of such families, frequently due to low levels of human capital, acts to increase income inequality in a region. Note that the coefficient on *FEMALE* is very stable, exhibiting very little change as *TIME* and then the regional dummy variables are added to the model.

The coefficient on the log of average family income (*LINCOME*) is also positive and significant. This result is similar to several other studies that found that income inequality begins to rise with higher incomes (see, for example, Garafalo and Fogarty, 1979). Levernier, Partridge, and Rickman (1998b) suggest that as market rewards for high-tech employment increase relative to jobs requiring lesser skills, the existence of a bimodal distribution of income could lead to greater income inequality. However, one must be cautious making conclusions concerning income because the coefficient on the variable is significant only when the regional dummy variables are added, implying that a stability issue exists.

The coefficient of the percent of the population that is under 18 (*UNDER18*) is negative and significant, while that for the percent of the population over 64 (*OVER64*) is not significant. The relationship for *UNDER18* also exhibits considerable stability as additional variables are added to the regression, indicating that the relationship is robust. This result is contrary to that of Levernier (1999), who found a positive and significant relationship for metropolitan counties for the percent of the population under 18. Perhaps having children spurs greater labor force effort, which results in more income, particularly at lower and middle income levels.

The coefficient on the female labor force participation rate (*FLFPR*) is also negative and significant at the 1 percent level. In addition, the coefficient estimates exhibit considerable stability as additional variables are added to the regression. As women enter the labor force in southeast Missouri, incomes of lower and middle income families are likely to be most affected, resulting in greater income equality.

The education variables in the model (*COLLEGE*, *HS*, *EDUC*) are generally insignificant except for the percent of the population 25 years of age or older that has a high school diploma, but no college degree. The relationship for *HS* is negative, indicating that income inequality falls as more of a county's population has a high school diploma. Southeast Missouri includes many counties where the population has relatively low rates of high school completion. However, high school completion rates have increased substantially over the past 10 to 20 years, and this has clearly led to greater income equality. The insignificance of *COLLEGE* may partially be a reflection of the low levels of college completion in the region. Note that the coefficient on *HS* is small and insignificant in the absence of the regional dummy variables. There is some question concerning the stability of this estimate; therefore, one must again be cautious in making conclusions concerning this variable.

In the absence of regional effects, coefficients on both the percentage of employment in the farm sector and the log of population are positive and significant at the 5 percent level. However, the fact that they lose significance when regional dummy variables are added indicates these two variables are related to regional effects and likely are not significant as explanatory variables.

The failure of the proportion of minorities (*BLACK*) to reach significance is an indication that either labor force discrimination is low in southeast Missouri or that *BLACK* is highly correlated with other social variables that do attain significance (such as female-headed families). Further research is necessary to ascertain the role, if any, of this variable in income inequality in the study region.

CONCLUSION

Recently, Federal Reserve Chairman, Alan Greenspan, in an appearance at a Joint Economic Committee hearing responded to a question by

Congressman Jack Reed that the observed significant divergence in the fortunes of different groups in the labor market "is not the type of thing which a capitalist society…can really accept without addressing" (Grier, 2005). The results of this study give way to some definite policy conclusions to address inequality. Income inequality in southeast Missouri can be reduced if the trend toward increased labor force participation of women continues. Policies, such as improved access to child care, that allow women to enter the labor force in yet greater numbers will reduce inequality. In addition, improved child-care choices should also help boost the incomes of female-headed families, though for these families, insufficient human capital may also be part of the equation. Therefore, job training or even high school completion policies (such as general equivalency diploma [GED programs]) could help to improve the economic fortunes of female-headed families and help to reduce income inequality. The significance of the percent of the population with a high school diploma in our regressions indicates that more than just female-headed families would benefit from high school completion policies.

The significance of the dependent population under 18 years of age in reducing income inequality, as indicated, may be due to the greater incentive to labor effort that having children can provide to families. Once again, access to adequate child care can help families with dependent children increase their labor effort.

It is apparent that there are similar forces at work here with respect to the significant variables in our model. Many of these forces revolve around access to the labor market, something that can be increased with better access to child care. In a recent study of child care in 20 counties in southeast Missouri, Birk et al. (2005) detailed the economic impact of the industry in the region. It is a large sector with a significant impact, and, as the results of this study show, it no doubt contributes to the reduction in income inequality in southeast Missouri.

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APPENDIX A

COUNTIES

The counties included in the planning commissions are as follows:

Bootheel:	Dunklin, Mississippi, New Madrid, Pemiscot, Scott, Stoddard
Lake of the Ozarks:	Camden, Laclede, Miller, Pulaski
Meramec:	Crawford, Dent, Maries, Phelps, Washington
Ozark Foothills:	Butler, Carter, Reynolds, Ripley, Wayne
South Central Ozarks:	Douglas, Howell, Oregon, Ozark, Shannon, Texas, Wright
Southeast Missouri:	Bollinger, Cape Girardeau, Iron, Madison, Perry, Ste. Genevieve, St. Francois

APPENDIX B

LIST OF VARIABLES

GINI	Gini coefficient (multiplied by 100)
OVER64	Percent of population over 64 years of age
UNDER18	Percent of population under 18 years of age
BLACK	Percent of population that is African-American
FEMALE	Percent of female-headed families
FLFPR	Female labor force participation rate
EMPGROW	Employment growth rate in previous decade
LINCOME	Log of average family income, constant 1982-84 dollars
LPOP	Log of population
MFG	Percent of employment in the manufacturing sector
FARM	Percent of employment in the farm sector
COLLEGE	Percent of population 25 or older with at least a Bachelor's degree
HIGH	Percent of population 25 or older with a high school diploma but no college degree
EDUC	Standard deviation of educational attainment
TIME	Dummy variable equal to 1 in 2000 and 0 in 1990
BOOTHEEL	Dummy variable equal to 1 for counties in Bootheel planning region
LAKEOZ	Dummy variable equal to 1 for counties in Lake of the Ozarks region
MERAMEC	Dummy variable equal to 1 for counties in Meramec region
OZFOOT	Dummy variable equal to 1 for counties in Ozark Foothills region
SCOZ	Dummy variable equal to 1 for counties in South Central Ozark region

A Spatial Analysis of Income Inequality in Arkansas at the County Level: Evidence from Tax and Commuting Data

John P. Shelnutt and Vincent W. Yao

In this paper, the authors examine income inequality at the county level in the state of Arkansas using data from individual tax returns. They find that the spatial pattern of inequality is positively correlated with economic growth. Therefore, the inverted-U hypothesis as it applies to regional income inequality is confirmed through cross-sectional analysis. This pattern can also be explained by many differences between metropolitan statistical areas (MSAs) and non-MSAs and cross-county commuting patterns. The important metropolitan area status–related variables include educational attainment, industrial composition, demographic distribution, and job-market condition. In an ordinary least-squares (OLS) model, these explanatory variables can explain most variations in the inequality. Commuting patterns also play an important role in explaining the inequality between job centers and fringe counties and between urban fringe and rural areas. The benefit of access to job centers is more significant in the MSAs than the micropolitan areas because of the quality and quantity of jobs available to commuters.

Federal Reserve Bank of St. Louis *Regional Economic Development*, 2005, *1*(1), pp. 52-65.

Since the 1990s or earlier, the debate over the relationship between income inequality and economic growth has intensified among economists. Traditional research on the topic has delivered a consistent message that the existence of inequality is detrimental to long-run economic growth. For instance, Alesina and Rodrik (1994) explain that when there is sizable inequality in a geographic area, the median voters will be poor. As a result, the political pressure from voters can direct government spending toward income redistribution and thus hurt investment and long-run economic growth. More recently, however, some studies have challenged this conventional wisdom. For instance, Forbes (2000) suggests that, in the short and medium term, an increase in a country's level of income inequality has a significant positive relationship with subsequent economic growth. Regardless of these disagreements, income inequality does have a relationship with economic growth that is coincidental, if not causal. Thus, more attention should be paid to this issue in policy implementation at the regional level, where economic development appears to be the most dominant objective.

The study of regional income inequality in the United States has thus far remained at the state level. Data are limited when analysis expands to a more detailed level. Williamson (1965) shows that regional inequality at the state level also follows the inverted-U curve found in the international pattern, increasing in early stages of economic development and decreasing in later

John P. Shelnutt is the chief economist at the Arkansas Department of Finance and Administration and Vincent W. Yao is a senior research economist at the Institute for Economic Advancement, University of Arkansas at Little Rock. The authors thank Rachel Kluender and Nick Nelson for the excellent GIS support and Hong Liu for expert research assistance. The helpful comments from Anthony Pennington-Cross and Bruce Domazlicky, as well as editors, are also greatly appreciated.

stages. Fan and Casetti (1994) further find that regional inequality has also been changing over time—the changes being associated with economic growth, sectoral shifts, and global spatial restructuring. The exceptions are four states in the South: South Carolina, Alabama, Mississippi, and Arkansas have remained, since the 1960s, in the high-inequality category, regardless of the emergence of the Manufacturing Belt in the 1960s, services sectors in the 1980s, and global competition in the late 1980s and 1990s.

This study examines the inequality patterns within the state of Arkansas, a state with one of the highest inequality rates since the 1960s. Most of the regional inequality literature thus far has focused on inequality across states, using U.S. census data on income at the county level. Inequality at the state level is usually measured by percentile ratio (e.g., Wheeler, 2004), Gini coefficient (e.g., Forbes, 2000), or Theil's T (e.g., Janikas and Rey, 2004). Amos (1988) has shown that the relationship between inequality and other factors (such as growth) at the state level is different from that at the county level. Moreover, a spatial analysis of income inequality within a state has far more interesting implications to the development of public policy and economic growth. Unfortunately, none of these inequality measures are readily available at the county level.

Using data from individual tax returns, the authors were able to construct an alternative measure of income inequality at the county level: the ratio of households in the top 25 percent of tax brackets to those in the bottom 25 percent of tax brackets. The analysis makes use of annual state tax filing data provided by the Arkansas Department of Finance and Administration. This data set was chosen for the current study and follow-up investigation because of its time-series availability and robust coverage of household information at the county level. Although decennial U.S. Census information is widely used in inequality studies, it was not used in this analysis. Instead, we have observed considerable fluctuation in the time-series information and change in income inequality. Our analysis here focuses on cross-sectional county patterns and factor relationships in the 2003 data. Table 1 includes the number of Arkansas tax returns (households) filed in each tax bracket (net taxable income). There are a total of 65 reported tax brackets, ranging from "under zero" to "$500,000 & over." The bottom 25 percent of tax brackets includes those households with net taxable income "under zero" through the "$15,000 to $15,999" bracket (or less than $16,000). The top 25 percent of brackets includes those households with net taxable income that is more than $49,000. To be sure, measurement errors are possible; current tax shelter programs could distort the relationship between a household's designated tax bracket and its actual income level. However, an assumption can be made behind the measurement: When filing a tax return, a rich household can downgrade itself by a few tax classes using some shelter provisions; however, it is unlikely to fall into the classification of a "poor household," because of alternative minimum tax (AMT) provisions and the limited nature of state tax exemptions. Therefore, the numbers of households in top 25 percent and bottom 25 percent of tax brackets provide relatively valid measures of the numbers of rich and poor households. A percentile analysis of the number of tax returns in each tax bracket is less effective than one based on income levels, such as that used in Wheeler (2004).

With more rich people and fewer poor people, the income gap narrows; with fewer rich people and more poor people, the income gap widens. Across counties, the higher ratio of rich people to poor indicates a higher income level and a higher level of inequality, and vice versa. Figure 1 plots the spatial distribution of income inequality of the 75 counties in Arkansas in 2003, which is the dependent variable we used in various models. There are six counties whose ratio is higher than 1.0: Benton in the northwest and Saline, Faulkner, Lonoke, Pulaski, and Grant in the central region. Other counties with relatively high inequality include Washington in the northwest, Ashley and Cleveland in the southeast, and Craighead in the northeast. Counties in the north and the west generally have low income inequality as well as a low income level.

Other data used in the study are obtained from the U.S. Census 2000. All the abbreviations

Table 1

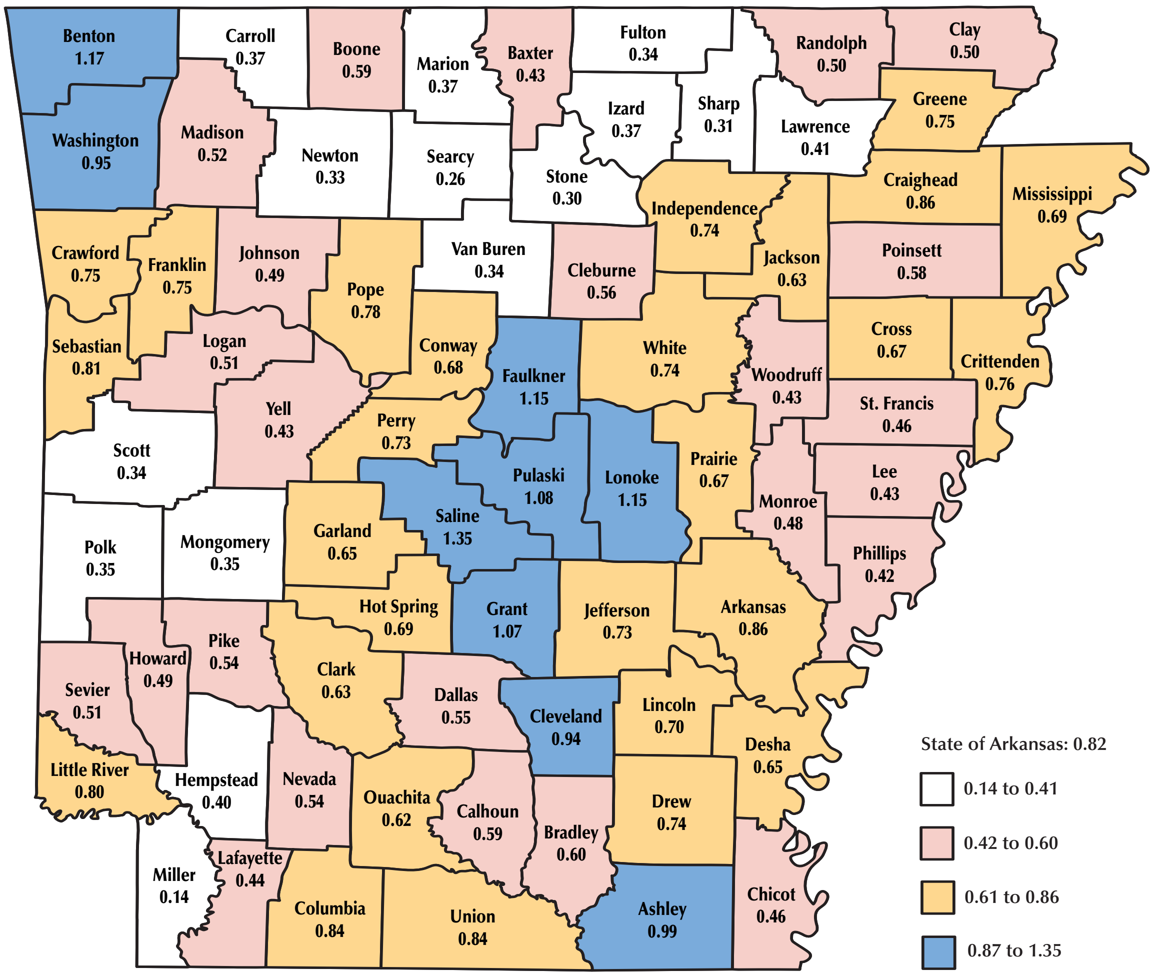
Arkansas Tax Data in 2003: State Total

Net taxable income ($)	Number of returns (household)
Bottom 25% of the tax brackets	
Under zero	1,678
0 TO 999	8,896
1,000 TO 1,999	5,202
2,000 TO 2,999	5,451
3,000 TO 3,999	5,905
4,000 TO 4,999	5,890
5,000 TO 5,999	6,084
6,000 TO 6,999	6,403
7,000 TO 7,999	9,039
8,000 TO 8,999	18,254
9,000 TO 9,999	21,979
10,000 TO 10,999	23,894
11,000 TO 11,999	18,320
12,000 TO 12,999	22,865
13,000 TO 13,999	24,072
14,000 TO 14,999	28,801
15,000 TO 15,999	31,222
16,000 TO 16,999	24,412
17,000 TO 17,999	21,126
18,000 TO 18,999	20,411
19,000 TO 19,999	19,755
20,000 TO 20,999	18,948
21,000 TO 21,999	17,962
22,000 TO 22,999	17,118
23,000 TO 23,999	16,476
24,000 TO 24,999	16,280
25,000 TO 25,999	15,206
26,000 TO 26,999	14,895
27,000 TO 27,999	14,049
28,000 TO 28,999	13,376
29,000 TO 29,999	12,862
30,000 TO 30,999	12,223
31,000 TO 31,999	11,722
32,000 TO 32,999	11,241
33,000 TO 33,999	10,502
34,000 TO 34,999	10,271
35,000 TO 35,999	10,108
36,000 TO 36,999	9,589
37,000 TO 37,999	9,260
38,000 TO 38,999	8,945
39,000 TO 39,999	8,498
40,000 TO 40,999	8,468
41,000 TO 41,999	7,967
42,000 TO 42,999	7,662
43,000 TO 43,999	7,658
44,000 TO 44,999	7,167
45,000 TO 45,999	7,078
46,000 TO 46,999	6,760
47,000 TO 47,999	6,693
48,000 TO 48,999	6,412
Top 25% of the tax brackets	
49,000 TO 49,999	6,420
50,000 TO 54,999	28,026
55,000 TO 59,999	24,262
60,000 TO 64,999	20,002
65,000 TO 69,999	16,364
70,000 TO 74,999	13,630
75,000 TO 79,999	11,132
80,000 TO 84,999	9,231
85,000 TO 89,999	7,473
90,000 TO 94,999	6,064
95,000 TO 99,999	5,078
100,000 TO 149,999	23,785
150,000 TO 199,999	7,890
200,000 TO 249,999	3,929
250,000 TO 499,999	6,475
500,000 & Over	4,640

SOURCE: Arkansas Department of Finance and Administration, 2004.

Figure 1

Ratio of Households in the Top 25 Percent to the Bottom 25 Percent of Tax Brackets, 2003



SOURCE: Arkansas Department of Finance and Administration, 2004.

are explained in Table 2. In the remainder of this paper, the inequality differentials across counties are explained. The second section tests the hypothesis of regional convergence in Arkansas by analyzing the correlation between the inequality ratio and economic growth. The third section explains the income differential associated with the metropolitan statistical area (MSA) status. The explanatory power of MSA/non-MSA differences is also correlated with educational attainment, job-market condition, and industrial composition. The forth section explores the supplemental contribution of commuting patterns to the inequality distribution. The last section provides a conclusion and summary of the need for further research.

ECONOMIC GROWTH

Following Kuznet (1955), the literature implicitly assumes that income inequality is a consequence of economic growth, as implied by the inverted-U hypothesis. There is disagreement, though, about whether this relationship is positive

Table 2

Data Description

Variable	Description
Inequality	Ratio of households in the top 25% to the bottom 25% of the tax brackets, 2003
Growth	Annualized growth rate of total nonfarm personal income, 1993-2003
Income	Median household Income
Job	Job-market condition measured by employment/population
Education	Percentage of the population with a Bachelor's degree or higher
WorkingAge	Percentage of the population that is 25 to 44 years old
Industry	Share of finance, insurance, and real estate (FIRE) and other knowledge-based industries
MSA dummy	1 if MSA, 0 otherwise

SOURCE: The inequality measure is constructed from tax data provided by the Arkansas Department Finance and Administration; all the other variables are calculated from U.S. Census 2000.

or negative. If there is a negative relationship between the two, inequality could eventually be minimized by economic development efforts, as claimed by the conventional wisdom. If it is positive, the widening income gap might suggest the potential for subsequent economic growth, as implied by "Dr. Inequality" (Forbes, 2000).

Our model specification is consistent with the inequality literature. However, unlike the growth model in the literature, our model is designed to explain the spatial distribution of income inequality. Therefore, the dependent variable is inequality and growth becomes the regressor:

$$Inequality_i = C + \beta_1 Growth_i + u_i, \tag{1}$$

where i represents each county, *Inequality* is measured by the ratio of households in the top 25 percent to the bottom 25 percent of tax brackets in 2003, C represents the constant term, *Growth* is the annualized growth rate of total nonfarm personal income from 1993 to 2003, and u is the error term. Because the inequality measure is constructed only for the year 2003, the analysis in this paper is cross-sectional. The OLS results are reported in Table 3, where both coefficients are statistically significant. Those results show that inequality has a positive relationship with economic growth. A 5.32 increase in the inequality ratio corresponds to a 100 percent increase in the economic growth rate. Therefore, counties with higher economic growth rates, such as those in central and northwest Arkansas, tend to have higher income inequality; whereas counties with lower growth rates, such as Baxter and Izard in the north and Dallas and Bradley in the south, tend to have lower income inequality.

Table 3

Income Inequality and Economic Growth

Dependent variable: *Inequality*

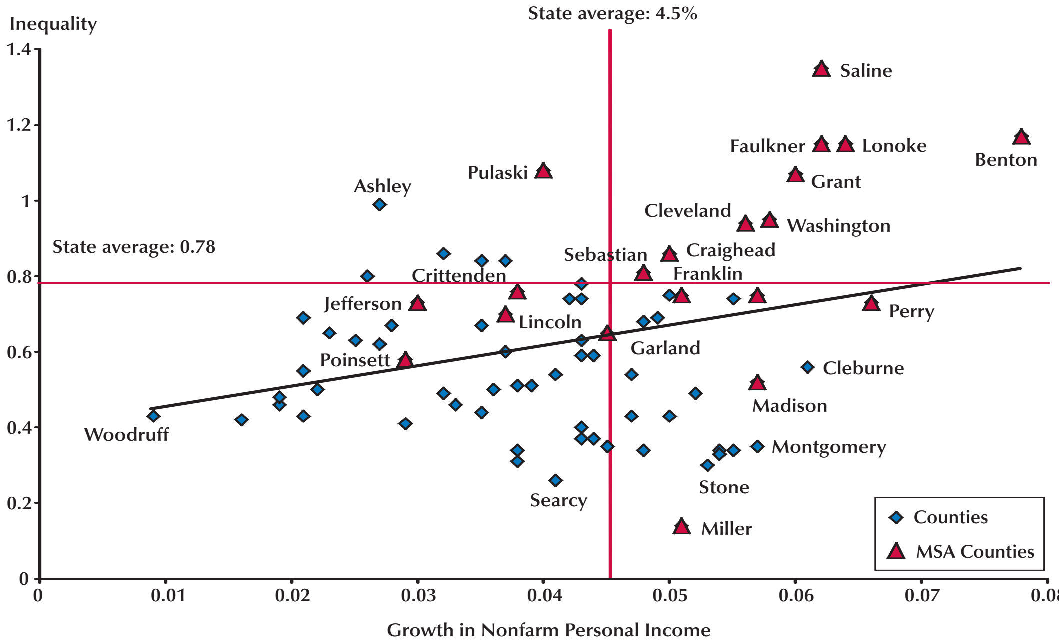
Variable	Coefficient	*t*-statistic
Constant	0.40	4.55
Growth	5.32	2.66
Adjusted R^2	0.08	

NOTE: All variables are statistically significant at the 5 percent level.

To further identify the performance of each county, Figure 2 plots the scatter diagram of the inequality ratios and economic growth rates of the 75 counties and the state averages. From 1993 to 2003, the state of Arkansas had an average growth rate for nonfarm income of 4.5 percent and an average inequality ratio of 0.78. Counties located in the northeast and southwest corners of the diagram display a positive relationship

Figure 2

Scatter Diagram of Inequality and Economic Growth



SOURCE: U.S. Bureau of the Census, U.S. Census 2000.

between inequality and growth (i.e., either high inequality and high growth or low inequality and low growth). There are nine counties that have both a higher inequality ratio and higher economic growth rate than the state averages: Benton, Saline, Lonoke, Faulkner, Grant, Cleveland, Washington, Sebastian, and Craighead. All of these are located in MSAs, where inequality generally accompanies rapid economic growth or significant gains from commuting. Among the counties located on the other end of the trend line, such as Woodruff and Searcy counties, low inequality is accompanied by lower economic growth.

Some counties also show a negative relationship between inequality and economic growth. (See the northwest and southeast corners of Figure 2.) Both Pulaski and Ashley counties have relatively high inequality (as shown in Figure 1), but they have a lower growth rate compared with other high-inequality counties. Pulaski County, where Little Rock is located, is the largest job center in the state. Ashley is home to many highly paid paper mill workers. Miller County, which is part of a border MSA (Texarkana) with Texas, has the lowest inequality ratio. Evidence of significant cross-border migration in Texarkana is associated with divergent tax treatment on income and usury lending effects on the Arkansas side of the border. Counties with higher inequality are most often located in MSAs, whereas those with relatively low inequality are most often located in non-MSAs. We therefore consider the MSA-related variables in the next section.

MSA OR NON-MSA?

Following the literature, an analysis of regional characteristics in the previous section has led us to describe inequality essentially as an outcome of growth processes. But inequality also has something to do with the strategic status of the region and its socioeconomic characteristics.

Fan and Casetti (1994) explain regional income inequality in the context of three phases of regional growth and industrial composition. Phase 1 involves initial advantages and agglomeration of activities associated with a "leading sector" and results in the formation of a core region of a state that is separate and distinct from peripheral areas. The formation of the state's core is accelerated by the movement of labor and capital from the periphery. A typical example in the United States was the emergence, consolidation, and widening of the Manufacturing Belt in the Northeastern and Midwestern states. Phase 2 is characterized by slower growth, stagnation, and the decline of areas within the main core and new growth in the former periphery. Phase 3 is driven by the spatial restructuring influenced by sectoral shifts and global competition.

Their theory was largely designed to assess inequality between states or nations, but a similar concept applies to inequality at the county level. Arkansas, as part of the national periphery, did not benefit to the same degree as other states did from the emergence of the Manufacturing Belt. Instead, Arkansas supplied labor to industrial states as part of its shift from an agriculture-based economy. However, as part of the polarization reversal in the 1960s and early 1970s, firms favored expansion into new locations in the periphery for their advantages such as a lower rate of unionization, lower labor and land costs, and an attractive climate. A good example of such growth and agglomeration is the emergence of the new MSA in northwest Arkansas, where Wal-Mart, Tyson Foods, and many trucking companies are based. Within the state, the difference between MSA and non-MSA, or urban and rural, is analogous to the core and periphery case, where MSAs attract most industries and jobs and thus have higher income inequality. Figure 3 maps the current metropolitan, micropolitan, and combined statistical areas in Arkansas. About 57 percent of the population resides in 20 counties of seven MSAs. For MSAs, the ratio of households in the top 25 percent to households in the bottom 25 percent of tax brackets is 0.93, whereas the ratio for non-MSAs and the state average are 0.58 and 0.78, respectively. The income inequality in MSAs is 60 percent higher than that in non-MSAs.

The above result is consistent with relationships Wheeler (2004) observed for seven states in the Federal Reserve's Eighth District. The difference between MSAs and non-MSAs also includes a wage premium in urban areas in general and a wage premium for highly educated individuals. This characteristic is evidence of the skill-biased technological changes that have taken place nationally. As with the well-known "New England Turnaround," the MSAs in Arkansas attract most capital flows and jobs, mainly due to the following: well-established infrastructure; emergence of knowledge-intensive industries; and availability of education, training, major health care systems, and centralized services. As shown in Table 4, which reflects data from the U.S. Census 2000, MSAs generally contain a higher percentage of well-educated individuals in the 25-to-44 age range and provide more job opportunities (employment/population). In MSAs, economic growth is faster and absolute income is higher than in non-MSAs. MSAs are both job centers and population centers for their regions. Not only are most people employed by businesses in these areas, but also jobs are concentrated in the faster-growing industries such as finance, insurance, and real estate (FIRE) and other professional scientific and business services.

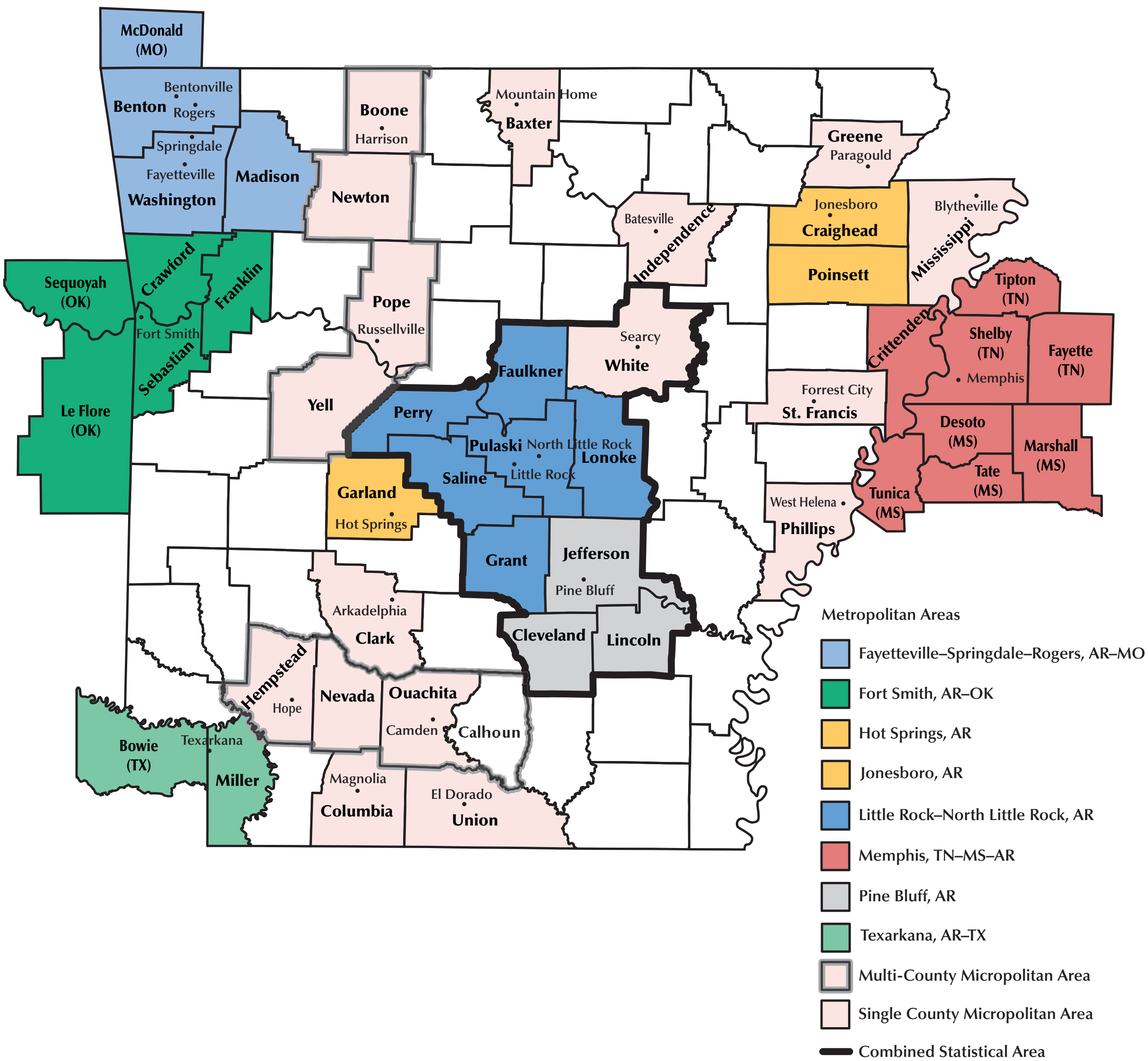
To quantify the contribution of these MSA-related variables, the following model was constructed:

(2)
$$Inequality_i = C + \beta_1 Job_i + \beta_2 Education_i + \beta_3 WorkingAge_i + \beta_4 Industry_i + u_i,$$

where i denotes county, C represents the constant, *Job* represents job-market condition measured by employment/population, *Education* is measured by percentage of Bachelor's degrees or higher, *WorkingAge* represents the share of persons aged 25 to 44 in the total population, and *Industry* is approximated by the share of FIRE and other knowledge-based industries. All of the above regressors are obtained from U.S. Census 2000 data, and the variables are explained in Table 2.

Figure 3

Metropolitan, Micropolitan, and Combined Statistical Areas in Arkansas (effective June 9, 2003)



SOURCE: U.S. Census Bureau and the Arkansas State Data Center.

Table 4

Differences Between MSAs and Non-MSAs

	MSA	Non-MSA	State average
Ratio of households by income (75th/25th percentile)	0.93	0.58	0.78
Growth	5.1%	3.8%	4.5%
Income	$39,681	$33,248	$35,071
Employment/population	57%	52%	53%
Percent of the population with a Bachelor's degree or higher	10%	8%	8%
Percent of agriculture, forestry, fishing, hunting, and mining	4%	7%	6%
Percent of FIRE and other knowledge-based industries	10%	7%	8%
Percent of the population that is 25 to 44 years old	29%	26%	27%

SOURCE: U.S. Census 2000.

The OLS regression results are reported in Table 5. Because there is a significant relationship between the MSA dummy variable and MSA-related variables, the former is excluded from the model to avoid multicollinearity. The included four regressors plus the constant term are all statistically significant at the 5 percent level. Together they account for 60 percent of the variations in income inequality. Although the constant term does not, the four explanatory variables all have a positive effect on income inequality. Counties tend to have a higher degree of inequality when they have job opportunities. Jobs are more likely to be created by fast-growing industries such as FIRE and other knowledge-related and concentrated industries. As stated above, growing sectors are more likely to be located in MSAs because they require better education and up-to-date skills as well as sophisticated infrastructure systems. Because of these skill-based needs, market forces favor individuals in the 25 to 44 age range. The returns to skilled workers result in greater separation of income groups. Thus, the income gap widens.

Because Table 4 suggests that there is a significant gap between the growth rates of MSAs and non-MSAs (5.1 percent versus 3.8 percent, respectively), we also report the correlation coefficients of these MSA-related variables in Table 6. All variables are correlated to a certain degree with one another, but not significantly. Therefore, each variable explains part of the inequality pattern across counties. The diagnostics of the regression do not suggest any misspecification.

Table 5

Income Inequality and MSA-Related Variables

Dependent variable: ***Inequality***

Variable	Coefficient	*t*-statistic
Constant	–1.22	–5.13
Job	0.96	2.39
Education	2.08	2.18
WorkingAge	3.60	3.73
Industry	2.44	1.97
Adjusted R^2	0.60	

NOTE: All variables are statistically significant at the 5 percent level.

IMPORTANCE OF COMMUTING PATTERNS

Apart from the factor as stated endowments in the MSAs, the working-age population living in other counties can still benefit by commuting to the job centers located in MSAs or micropolitan areas. Thus, commuting patterns and access to highway corridors are also important to income inequality. Figure 4 depicts the net gain or loss of

Table 6

Correlation Matrix of All MSA-Related Variables

	Inequality	*Growth*	Residuals from model (1)	*Job*	MSA dummy	*Working Age*	*Education*	*Industry*
Growth	0.30							
Residuals from model (1)	0.95							
Job	0.60	0.47	0.48					
MSA dummy	0.51	0.45	0.40	0.47				
WorkingAge	0.61	0.28	0.56	0.49	0.61			
Education	0.61	0.34	0.54	0.49	0.43	0.33		
Industry	0.63	0.38	0.55	0.45	0.57	0.44	0.72	

SOURCE: U.S. Census 2000.

workers commuting in or out of individual counties. By counting those counties with 7 percent or more of the workforce commuting-in, there are several major job centers in the state: Pulaski in the central region, Sebastian in the west, Boone in the north, Independence and Craighead in the northeast, Arkansas-Desha in the east, Union in the south, and Howard and Clark in the southwest. They can be divided into three types: in-state MSAs, cross-state MSAs, and stand-alone small job centers (or micropolitan areas).

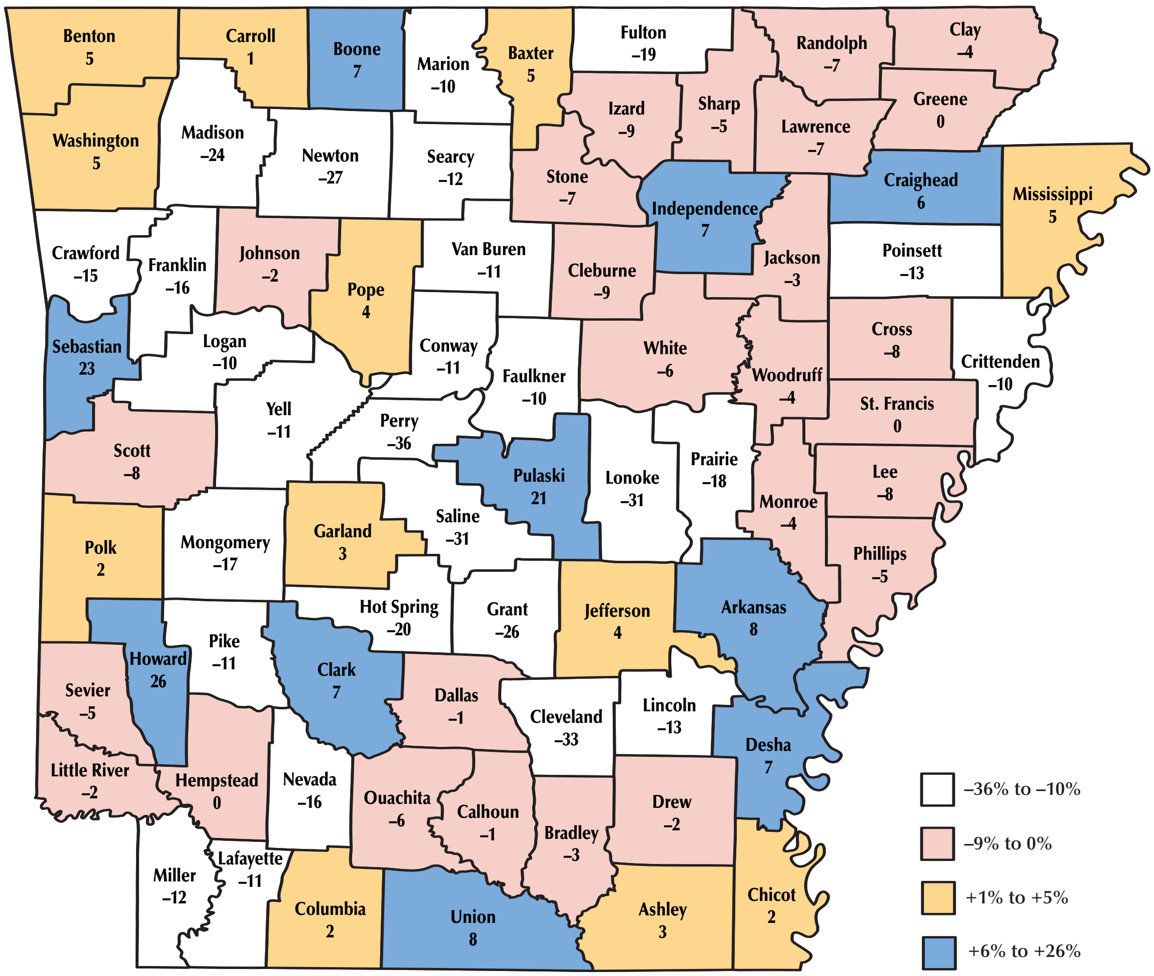
The largest in-state MSA is the Little Rock–North Little Rock area in central Arkansas, which includes Pulaski County in the urban center. As shown in Figure 4, the net gain/loss in commuter flow is 21 percent in Pulaski County and negative in all but one of the fringe counties. Individuals commute from the fringe counties to the Little Rock area to work. Consequently, all of these counties have high inequality because of better access to job markets, a large working-age population, and other advantages in the metropolitan area. In northwest Arkansas, Benton, Washington, and Carroll counties benefit from the presence of large corporations such as Wal-Mart, Tyson Foods, and J.B. Hunt Trucking. It is the most rapidly growing area in the state, and commuters from the fringe counties benefit from job creation and generally higher wages in these areas. Craighead County has become the urban center in northeast Arkansas because of Arkansas State University and other manufacturing firms located in Jonesboro. Jefferson County also attracts capital and workers because of the economic growth and development in Pine Bluff.

There are also three MSAs shared between Arkansas and other states. The MSA centered on Sebastian County is shared with Oklahoma and includes Fort Smith, the third largest city in the state. It thus attracts commuters and capital from the fringe counties in Arkansas and adjacent counties in Oklahoma. Crittenden County is located within an MSA shared with Memphis, Tennessee, and it benefits from the flow of commuters to the Tennessee side of the border. Miller County also benefits from the commuter flow to the Texas side of Texarkana.

There are several micropolitan areas within the state, which generally center on a single town as the economic driver in that area. For instance, Harrison (Boone County) attracts commuters because of the presence of a FedEx branch and several manufacturing firms in the furniture and wood products sectors. Batesville (Independence County) is home to several poultry processing and chemical industries. It also benefits from a fast-growing community college and a private four-year college. Hot Springs, a small MSA in Garland County, provides a major venue for tourism, convention and hospitality, and regional services.

Figure 4

Net Gain or Loss in Commuter Flow



SOURCE: U.S. Census 2000 and calculations by the authors.

Each of these micropolitan cases provides considerable services and retail capacity to surrounding rural counties. Notable examples of services include regional hospitals, small airports, and a variety of legal and financial services. Other job centers in non-MSA and non-micropolitan designations consist of major industrial facilities tied to forest resources; for example, large paper mill operations in rural counties provide high-wage jobs in sparsely populated areas—notably, in Ashley and Little River counties. Similar industrial location factors are noted in Union County.

Table 7 further organizes average income inequality levels by different statistical areas (MSAs and micropolitan areas) and commuting pattern. All 75 counties in the state of Arkansas are classified into one of five categories, with the number of counties of each category in parentheses:

- MSA urban center (7 counties): those with positive commuter flows in the MSA, including Pulaski, Washington, Benton, Sebastian, Jefferson, Craighead, and Garland;

Table 7

Inequality by Statistical Areas and Commuting Pattern

Statistical areas in Arkansas	Number of counties	Percentage of net commuter flow (percent)	Inequality
MSA	20	3	0.84
Urban job center	7	12	0.89
Urban fringe	13	–19	0.81
Micropolitan area	17	0.3	0.60
Micropolitan job center	7	6	0.69
Micropolitan fringe	10	–5	0.53
Rest of the state	38	–3	0.53
Entire state	75	0	0.78

SOURCE: U.S. Census 2000.

- MSA urban fringe (13 counties): those with negative commuter flows in the MSA;
- Micropolitan center (7 counties): those with positive commuter flows in the micropolitan areas, including Boone, Baxter, Pope, Independence, Clark, Columbia, and Union;
- Micropolitan fringe (10 counties): those with negative commuter flows in the micropolitan areas;
- Rest of the state (38 counties).

On average, 12 percent of the urban centers' working population commute in from fringe counties to work. About 20 percent of the working population in fringe counties commutes into job centers in an MSA. Overall, about 3 percent of the working population in MSA counties commutes from outside of the MSA, which implies that MSAs as a whole provide job opportunities to the rest of the state. Commuter flow is significantly lower in micropolitan areas. Only about 6 percent of the working population in job centers commutes in from fringe counties, and about 5 percent of those in fringe counties commutes out. Micropolitan areas, overall, have a net flow of commutes of 0.3 percent. The rest of the state has a net pattern of commuter out flow. About 3 percent of the state's population commutes to either MSAs or micropolitan areas to work. Overall, the labor market in Arkansas is self sustained, with 0 percent net commuter flow.

As shown in Table 7, the commuting pattern has a strong linkage with the inequality distribution and, in turn, economic growth. Households in urban centers usually have the highest income inequality, followed by urban fringe counties, micropolitan job centers, micropolitan fringe counties, and, finally, the rest of the state. In addition to the significant inequality gap between MSA and non-MSA counties, another gap exists between commute-in (job center) and commute-out (fringe) counties. The latter gap is wider in micropolitan areas than in MSAs. However, the average inequality ratio of micropolitan areas (0.60) is lower than the state average (0.78) and marginally exceeds the rest of the state (0.53). The inequality ratios are similar for micropolitan fringe counties and the rest of the state, although the former have a higher percentage of commuting-out. Therefore, commuting promotes income inequality and economic growth in urban areas more so than in rural areas. Larger-scale and higher-paying jobs are concentrated in urban centers, which allows more people to commute in and thereby benefits the MSA by, among other things, promoting relatively higher wages. These concentration and scale differences partly explain the high inequality in urban fringe relative to rural areas. On the other side, jobs in micropolitan areas are not compensated as well as those in urban areas and their scale is also not that large. Thus, access to jobs in these areas does not help change the income level, economic growth, or inequality of the micropolitan fringes relative to the rest of the state (0.53 vs. 0.53). There is another possible explanation for the similarity in the inequality level of micropolitan fringes and that of rest of the state: measurement error. Access to job centers or the possibility of commuting promotes economic growth and income improvement and eventually affects inequality. However, tax data may not reflect the difference because of the distortion caused by tax shelter programs.

CONCLUSIONS

Using the data from individual Arkansas tax returns, this study develops an indicative measure of income inequality at the county level. In the state of Arkansas, income is most unequally distributed in the northwest and central portions of the state and selected counties in the southern part of the state. This spatial pattern is positively correlated with economic growth. Counties differ in their inequality over the course of their economic development and inequality may decrease as the economy develops. The inequality pattern can also be explained by many factor differences between MSAs and non-MSAs, such as educational attainment, sector composition, demographic distribution, and job-market conditions. This paper also uses the data on commuting patterns to show that a fringe county can still benefit when its population commutes to a nearby job center. However, access to urban centers is more beneficial than access to micropolitan areas because the job quality in the latter is much lower.

This study provides an Occam's razor for policymakers by separating the intertwined issues of income inequality and growth at the regional level. Although more research is needed, the preliminary findings point to area growth and urban concentration as principal drivers for income inequality over time and in spatial distributions. The study also implies that concern about income inequality, or more likely the rate of change in inequality, would be better directed at addressing root factors rather than social engineering or punitive tax policy. Factor analysis provided in this study points to the need for more effective educational systems and occupational training as constructive ways to respond to the effects of regional growth on incomes and the spreading effects on household income distributions that stem from rising opportunity in a given job center.

In addition to improving the specificity and clarity of inputs to regional growth, the study identifies the need for greater coordination in transportation-system planning overall and economic development of rural or micropolitan areas. The same growth and inequality relationships observed in large urban centers are observed down to the micropolitan and county levels. This study's policy implications are not unlike those of other studies that promote public and private investment in programs and infrastructure for rural development. The difference here is that the prospect of rising income inequality should not be a deterrent to growth and development efforts.

Further research is needed to measure and test the role and significance of commuting on income inequality. Part of this examination will need to account for several data issues and data methodological issues. The current research has shown a statistically significant relationship between metropolitan and nonmetropolitan areas of the state. Lack of statistical significance among counties with an elevated commuting rate has not been fully examined. Issues of county-level data quality and definitional shifts of MSAs need to be accounted for, given the accretionary changes in MSA designation over time and the definitional role of commuting ties to the urban core and income dependency ratios. Explanatory models of income inequality may be inefficient when combining MSA variables with non-MSA variables for commuting-dependent counties.

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Cost of Government Services: Trends and Comparisons for Kentucky and Its Neighboring States

William H. Hoyt, John E. Garen, and Anna L. Stewart

The authors examine expenditures for a variety of government functions for Kentucky and its neighbors (Illinois, Indiana, Missouri, Ohio, Tennessee, Virginia, and West Virginia) for 1992, 1997, and 2002. While per capita spending provides some gauge of the efficacy of public service provision, population may inadequately measure the client base or determinant of costs. To address this problem, they control for other factors that may influence expenditures, including population, age, and demographics. They believe this extensive quantification of costs and the comparison of these costs among states represent a unique effort in providing important information about service production for state governments. Although the authors do not offer conclusions regarding the efficacy of provision of public services, this study can aid state governments in their assessment of services.

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Government "waste" or occasionally even fraud has often been the subject of public concerns, political rhetoric, and investigative reports in the media. Yet, despite frequent overtures by elected officials about eliminating waste (or at least reducing costs) and occasional examples of claimed reductions in costs or elimination of waste, there have been very few examples in the popular press or even scholarly work of attempts to compare costs among governments. While there are legitimate concerns about how to interpret simple cost comparisons, such as those made here, it is still somewhat surprising that they are not made more often, given the amount of attention paid to government costs and relative taxation.

Here, our purpose is to make relatively simple comparisons of the costs of government services, both state and local, among the Commonwealth of Kentucky and its neighbors: Illinois, Indiana, Missouri, Ohio, Tennessee, Virginia, and West Virginia. Because this study was initially done for the Governor's Office of the Commonwealth of Kentucky, much of the discussion focuses on the costs of a variety of government services within Kentucky, for the years 1992, 1997, and 2002. In addition, employment and salaries in government services are also examined. For most of the services and government functions, cost comparisons are made on a per capita basis in 2002 dollars. Employment is also adjusted to reflect differences in population. Salary comparisons are adjusted for inflation and in some cases also adjusted to reflect differences in private earnings among the states.

While we think that this study can provide useful information for evaluating the relative efficacy of public service provision, we do not intend to imply that differences in costs by themselves, particularly when measured on a per capita basis, imply differences in the performance or

William H. Hoyt is a Gatton Endowed Professor of Economics, professor of public policy, and interim co-director for the Center for Business and Economic Research (CBER); John E. Garen is a Gatton Endowed Professor of Economics and interim co-director for the CBER; and Anna L. Stewart is an economic analyst at the CBER, Department of Economics, University of Kentucky.

efficiency in the provision of government services. For some services, population may not be a very accurate measure of the client base or determinant of costs. For a number of government functions, we use alternative measures as a base for costs. For example, education costs are on a per student basis and highway costs are on a mileage basis. While we believe that these alternative bases for costs more accurately reflect the determinants of costs, they, too, fail to reflect differences in the quality or extent of services.

In addition to providing data that indicate both trends and differences in the costs of government services and activities, we also provide some estimates of "cost" or expenditure functions for total state and local expenditures, administrative expenditures, and primary and secondary educational expenditures. We have three primary objectives in estimating these relationships. Our first objective is to determine how much of the expenditures within a state cannot be explained by controlling for factors that might affect either the cost or quality of services within the state (including state population and the demographic composition of its population). Second, we use a fixed-effect model to estimate a state fixed effect for each state so that we may better understand some of the reasons expenditures across states may vary. To do this, we decompose the source of variation in predicted costs using the coefficients obtained in our estimation. Finally, we estimate and depict the impact of population on costs, that is, the existence of economies or diseconomies of scale primarily to understand and explain differences in costs among states, but also to better understand what might constitute the "ideal" population of a jurisdiction, state or locality, for the purposes of public service provision.

While we estimate a relationship between (i) expenditures and (ii) factors likely to influence expenditures that might be referred to as a "cost" function, we are reluctant to ascribe that nomenclature to it. Numerous reasons for differences in spending are possible. One limitation in our analysis is the difficulty in reliably measuring "output" of a government service or function or the quality with which it is provided. When possible, we do attempt to measure the number of customers or clients (vehicle miles for highways and students for education, for example), but even these measures do not control for differences in the quality of services.

Despite these qualifications about the measurement of both the quantity and quality of government services, we believe that the measurement of costs among state governments in this study represents a unique effort and provides important information about service production. Again, although the evidence presented in this study is not, by itself, conclusive regarding the efficacy of provision of public services, we believe it can direct state governments in assessing particular services more thoroughly. While the more qualitative approach used in typical performance evaluation studies has value, we believe that our focus on costs complements the approach in these other studies of assessing quality in performing a service.

The study is designed to focus on state government services, but there is significant variation among the states to which we compare Kentucky in the responsibilities of state and local governments. Kentucky, along with West Virginia, has the greatest share of state and local spending that is financed by the state. Therefore, for most of the services we examined, we believed it important to examine both state and combined state and local spending and employment. In addition, even if the spending is not done at the state level, the state is frequently the financer of these expenditures, particularly for Kentucky.

In our sample of states, we find that less populous states and states with more centralized spending have higher per capita expenditure. Regarding particular government functions, no clear patterns emerge for central administration expenditure and employment, though low-wage states, especially Kentucky and West Virginia, tend to have high central administration salaries. Regarding primary and secondary education spending, with the exception of West Virginia, spending per student is higher for states with higher income and larger populations. All states experienced a reduction in the student-teacher ratio, but (with the exception of Missouri) also a reduction in the student-administrator ratio. The least populous states—West Virginia and Kentucky—have the highest per capita spending on highways. This does not hold for spending per road mile, however. West Virginia stands out as exceptionally high in employment in highway provision.

Our multivariate analysis reveals some interesting findings, too. There are economies of scale—more populous states have less spending per capita. States with more centralized spending have more state and local total spending and higher-wage states have greater spending. States with a greater population per municipality and a higher poverty rate have lower spending. Controlling for more covariates tends to raise the estimated scale economy. The state fixed effects change substantially after controlling for the covariates. More populous states now tend to have higher expenditure. The results of the multivariate analysis for central administration spending tend to mirror the findings for total expenditure.

For primary and secondary educational expenditures, economies of scale are strong for students per district, but less so for students per school. Higher-wage states have higher expenditure per student. Measures of student performance (i.e., test scores) have little relationship to spending. Control for covariates alters the estimated differences between states, but the ranking does not change much.

In the following section we provide some data on the demography of Kentucky and its neighboring states, as well as some information about the economic structure of these states. These data are from the 2000 (and 1990) Census of Population and Housing. We then report on aggregate government spending and employment without regard to government functions or services. We then report on government spending, employment, and earnings by government functions, including central administration, financial administration, primary and secondary education, and highways and roadways. The penultimate section reports the findings of our regression analysis.

SOME BASIC FACTS ABOUT THE KENTUCKY POPULATION

Table 1 contains data from the 2000 U.S. Census of Population[1] on characteristics of Kentucky's and its neighbors' populations. Information on employment is obtained from the Regional Economic Information System (REIS).[2] As Table 1, Panel A, shows, Kentucky is the second smallest state (in population) in this group of states and is the second most rural. It is ranked eighth when compared with neighboring states as well as the United States overall in the percentage of its population that is African-American. It is also ranked eighth in the percentage of its population that is Hispanic. The percentage of households with children under 18 years of age in Kentucky is very similar to its neighboring states and the U.S. average; it ranks relatively low in the percentage of households over 65 years of age.

Table 1, Panel B, provides U.S. Census data on income, earnings, and employment. Again, Kentucky's income (both median family and per capita) and earnings (for ages 16 and older) are above only West Virginia's levels and only West Virginia has a higher poverty rate. In 2000, Kentucky's unemployment rate (5.7 percent) was approximately the same as that in the United States (5.8 percent) and in the middle of the range of these states; yet, it had the lowest employment rate, that is, the percentage of its adult population (ages 16 and older) employed. A relatively high percentage of respondents to the survey in Kentucky reported themselves as disabled, meaning that a disability impairs their ability to be employed or function in their job if employed.

AGGREGATE AND CURRENT GOVERNMENT SPENDING

Before considering spending on each of the several government functions in detail, we first provide some recent data on aggregate spending

1 These data are available electronically from the U.S. Census Bureau, www.census.gov. Data in Tables 1A and 1B are from Census 2000, with the exception of the estimate of populations for 2003, which are also available at the Census website and are obtained from estimates made by the Bureau of Economic Activity (BEA).

2 The REIS is produced by the Bureau of Economic Activity using data obtained from County Business Patterns: www.bea.doc.gov/bea/regional/data.htm.

Table 1

Population Characteristics of Kentucky and Its Neighbors

A. Population and Population Composition, 2000

	Population		Urban		White		African-American		Hispanic		Households with children under 18		Households over 65	
State	# (2003)	Rank	%	Rank	%	Rank	%	Rank	%	Rank	%	Rank	%	Rank
Kentucky	4,117,827	7	55.8	8	90.1	2	7.3	8	1.5	8	33.0	5	31.6	7
United States	290,809,777		79.2	2	75.2	7	12.3	4	13.7	1	33.5	2	33.1	4
Illinois	12,653,544	1	87.8	1	73.5	8	15.1	3	12.3	2	33.6	1	32.6	5
Indiana	6,195,643	4	70.8	5	87.5	3	8.4	7	3.5	4	33.4	3	32.2	6
Missouri	5,704,484	6	69.4	6	84.9	5	11.2	6	2.1	6	32.4	6	34.4	2
Ohio	11,435,798	2	77.4	3	85.0	4	11.5	5	1.9	7	32.2	8	33.9	3
Tennessee	5,841,748	5	63.6	7	80.2	6	16.4	2	2.2	5	32.2	7	31.5	8
Virginia	7,386,330	3	73.0	4	72.3	9	19.6	1	4.7	3	33.2	4	29.3	9
West Virginia	1,810,354	8	46.1	9	95.0	1	3.2	9	0.7	9	29.3	9	37.5	1

B. Income and Employment Measures, 2000

	Median household income		Income per capita		Income below poverty level		Median earnings		Unemployed		Employed		Disabled	
State	2000$	Rank	2000$	Rank	%	Rank	2000$	Rank	%	Rank	%	Rank	%	Rank
Kentucky	33,672	7	17,819	7	15.8	2	20,951	7	5.7	4	44.5	8	9.9	3
United States	—		21,067		12.9		—		5.8		45.8		8.2	
Illinois	46,590	2	22,760	2	10.7	6	25,890	1	6.0	2	47.0	5	8.8	6
Indiana	41,567	3	20,076	4	9.5	9	23,229	4	4.9	8	48.8	1	7.2	8
Missouri	37,934	5	19,618	5	11.7	5	21,751	5	5.3	6	47.5	4	10.9	2
Ohio	40,956	4	20,694	3	10.6	7	23,949	3	5.0	7	47.6	3	9.6	4
Tennessee	36,360	6	19,120	6	13.5	3	21,700	6	5.4	5	46.6	6	11.2	1
Virginia	46,677	1	23,506	1	9.6	8	25,357	2	4.1	9	48.2	2	9.5	5
West Virginia	29,696	8	16,322	8	17.9	1	19,159	8	7.3	1	40.5	9	6.4	9

SOURCE: Census 2000 (www.census.gov).

Table 2

State and Local Total Expenditures Per Capita (2002$), Selected Years

	Per capita, 2002$			Rank			
	1992	1997	2002	1992	1997	2002	Annual % change
Kentucky	4,697	5,170	6,073	7	7	6	2.6
United States	5,865	6,217	7,125	1	1	1	2.0
Illinois	5,230	5,843	6,944	3	2	3	2.9
Indiana	4,568	4,966	5,896	8	8	8	2.6
Missouri	4,255	4,838	5,827	9	9	9	3.2
Ohio	5,357	5,746	7,010	2	4	2	2.7
Tennessee	5,112	5,775	6,328	4	3	5	2.2
Virginia	4,797	5,344	5,994	6	6	7	2.3
West Virginia	4,896	5,564	6,609	5	5	4	3.0

Table 3

State Share of State and Local Expenditures, 2002, by Function (percent)

	Higher education	Primary and secondary education	Public welfare	Health	Highways	Correction	Parks and recreation	Financial administration	Judicial and legal services
Kentucky	100	67	99	51	81	65	47	74	82
United States	84		85	53	61	68	16	55	46
Illinois	68	37	96	81	45	72	7	47	28
Indiana	100	55	89	73	64	74	11	51	30
Missouri	80	39	97	71	59	76	9	55	50
Ohio	92	49	80	30	54	77	11	54	17
Tennessee	100	48	98	75	64	58	26	38	47
Virginia	97	44	79	47	82	69	13	58	46
West Virginia	99	68	100	71	94	85	53	75	68

in Kentucky and its neighboring states.[3] In addition, we offer data suggesting how responsibilities for the revenue collection and the provision of government functions (expenditures) often differ significantly among states. State and local total expenditures per capita are shown in Table 2. While Kentucky ranks second for state spending per capita among these states, for combined state and local expenditure, Kentucky ranks sixth, reflecting more centralized expenditures.

As shown in Table 3, the share of state spending in total state and local spending is disaggregated by government function. For some functions, states are very similar in their allocation of spending between state and local governments. These are general functions performed exclusively by

[3] In this section and in the following tables, data on both state and local government spending and employment, unless otherwise indicated, are obtained from the U.S. Census Bureau surveys of state governments (U.S. Census Bureau Governments Division Annual Survey of Government Finances and Annual Survey of Government Employment), which were used to obtain figures (estimates) of government finances and employment in years in which a census is not undertaken (www.census.gov/govs/www/index.html).

Table 4

Mean Wage and Relative Wage (May 2003)

State	Wage	Rank	Relative to Kentucky	Relative to U.S.
Kentucky	15.15	8	1.00	0.86
United States	17.70	3	1.17	1.00
Illinois	17.95	1	1.18	1.01
Indiana	15.90	6	1.05	0.90
Missouri	16.23	5	1.07	0.92
Ohio	16.77	4	1.11	0.95
Tennessee	15.34	7	1.01	0.87
Virginia	17.76	2	1.17	1.00
West Virginia	14.20	9	0.94	0.80

state governments, such as social insurance and public welfare. With the exception of Illinois and Missouri, public higher education is primarily financed by state governments. Kentucky bears a much higher share of expenditures on highways, parks and recreation, and primary and secondary education than its neighboring states and the U.S. average. The same is true for financial, judicial, and legal administration. Only in health and corrections is Kentucky's state share below the national average, and, in these cases, it is only slightly below.

The significant differences in how spending is allocated between state and local governments among our group of states suggests that for much of our analysis the examination of state and local expenditures, rather than only state or only local, is appropriate.

Meaningful comparison of expenditures over time requires adjusting for changes in the base population—or, for some government goods or services, some measure of the good produced or population being served. For this reason we generally report expenditures on a per capita basis. In addition, changes in prices need to be accounted for when comparing expenditures over time. All expenditures here are reported in 2002 dollars, meaning that expenditures in early years (1992, 1997) are inflated to 2002 values using the consumer price index (CPI) produced by the U.S. Department of Labor, Bureau of Labor Statistics.

In addition to examining expenditures and employment, we also report trends and comparisons in salaries for the government functions. To make meaningful comparisons among the states and over time, we adjust the reported salaries in two ways. First, salaries are adjusted for inflation and reported in 2002 dollars, as is done with expenditures using the CPI. Second, we adjust for differences in the general level of salaries and wages among the states. Specifically, we create a wage index, reported in Table 4, to adjust for differences in the general level of wages and salaries among states. Thus, if a state has higher earnings in the private sector, salaries in the public sector will be deflated to reflect the higher private sector compensation in that state. As Table 4 shows, private sector workers in Illinois are paid, on average, 18 percent more than workers in Kentucky, so we would expect public sector employees to be paid more in Illinois as well. As the table shows, of the neighboring states, only West Virginia has lower wages on average.

The focus of Tables 5 and 6 is the salaries of state employees. Table 5 reports the average monthly salary of state employees adjusted for inflation but not adjusted for geographical differences in salaries. For all three years reported, Kentucky is ranked in the middle (fifth or sixth) in salaries, with average salary being almost $400 per month less than the U.S. average. However, when salaries are indexed based on differences

Table 5

Salaries, Average for All State Employees, Adjusted for Inflation

	Monthly salary, 2002$			Rank			
	1992	1997	2002	1992	1997	2002	Annual % change
Kentucky	2,797	2,873	3,115	6	5	5	1.09
United States	3,259	3,209	3,514	1	3	2	0.75
Illinois	3,257	3,349	3,583	2	1	1	0.96
Indiana	3,070	2,861	3,002	4	6	6	–0.22
Missouri	2,653	2,566	2,739	8	9	9	0.32
Ohio	3,225	3,249	3,419	3	2	3	0.58
Tennessee	2,708	2,712	2,865	7	7	7	0.57
Virginia	2,848	2,905	3,286	5	4	4	1.44
West Virginia	2,477	2,609	2,841	9	8	8	1.38

Table 6

Salaries, Average for All State Employees, Indexed and Adjusted for Inflation

	Monthly salary, indexed			Rank			
	1992	1997	2002	1992	1997	2002	Annual % change
Kentucky	2,797	2,873	3,115	3	2	1	1.09
United States	2,790	2,747	3,008	4	5	5	0.75
Illinois	2,749	2,826	3,024	5	3	4	0.96
Indiana	2,925	2,726	2,861	1	6	6	–0.22
Missouri	2,477	2,396	2,557	8	9	9	0.32
Ohio	2,914	2,935	3,088	2	1	2	0.58
Tennessee	2,674	2,678	2,829	6	7	7	0.57
Virginia	2,430	2,478	2,803	9	8	8	1.44
West Virginia	2,643	2,783	3,031	7	4	3	1.38

in mean wages, intended to reflect differences in local labor markets, the rankings change dramatically. Indexing for these differences in average state wages leads to Kentucky having the highest indexed salary among its neighbors in 2002. This finding indicates that while wages, both private and public, are on average 17 percent lower in Kentucky than the entire United States, the difference in salaries for state employees in Kentucky is not nearly this great, being only about 11.4 percent lower than the U.S. average. In determining an appropriate comparison for salaries adjusted only for inflation or salaries adjusted for inflation and general differences in salaries across the states, the nature and extent of the labor market for the state employee must be considered. For some occupations, the labor market is national or at least regional; for these occupations, local market conditions are not relevant and comparisons based on salaries should not be adjusted for geographical differences in wages. If, instead, state employees in an occupation are hired from local labor markets and tend to search within the state rather than the region or state, then the

Table 7

Salaries, Average for All State and Local Employees, Indexed and Adjusted for Inflation

	Monthly salary, 2002$			Rank			
	1992	1997	2002	1992	1997	2002	Annual % change
Kentucky	2,853	3,122	2,648	3	4	6	–0.74
United States	2,943	3,203	2,780	2	2	3	–0.57
Illinois	2,824	3,194	2,677	5	3	5	–0.53
Indiana	2,849	3,069	2,746	4	6	4	–0.37
Missouri	2,701	2,918	2,527	9	9	8	–0.66
Ohio	2,960	3,287	2,889	1	1	2	–0.24
Tennessee	2,722	2,937	2,614	7	8	7	–0.40
Virginia	2,707	3,009	2,504	8	7	9	–0.78
West Virginia	2,740	3,075	2,926	6	5	1	0.66

Table 8

State and Local Government Employment Per 1,000 Residents, Selected Years

	Employment per 1,000			Rank			
	1992	1997	2002	1992	1997	2002	Annual % change
Kentucky	52.47	53.14	56.25	3	4	2	0.70
United States	51.68	53.60	54.29	4	3	4	0.49
Illinois	47.51	50.35	51.08	9	8	9	0.73
Indiana	52.48	52.75	52.81	2	5	7	0.06
Missouri	47.64	54.10	55.12	8	1	3	1.47
Ohio	48.29	50.24	53.37	7	9	5	1.01
Tennessee	49.68	52.01	52.91	6	6	6	0.63
Virginia	54.60	53.81	56.40	1	2	1	0.33
West Virginia	50.31	50.70	51.61	5	7	8	0.26

salaries are adjusted for differences in mean wages in the state.

Table 7 reports the average salary, indexed and adjusted for inflation, for all state and local employees. In contrast to indexed salaries for state employees only, indexed salaries aggregated to include local employees are not particularly high. This, of course, suggests that local employee salaries must be quite low relative to those in other states. The ranking for Kentucky fell from third in 1992 to sixth in 2002, with average real salaries falling by an annual average of –0.74 percent, the biggest decrease except for salaries in Virginia. In contrast, real state salaries have risen 1.09 percent per annum, well above the national average of 0.75 percent.

Unlike salary comparisons, when state and local employment is combined (Table 8), Kentucky's level still remains very high, with 56.25 state and local employees per 1,000 residents in 2002. This is second only to Virginia's rate of 56.40 and is above the 2002 U.S. average of 54.29. Again, the differences in the distribution of government services between state and local governments are similar to the differences between the states in state and local employment; the latter differences are much smaller than those found when considering only state government employment. In contrast to state government employment alone, state and local government employment has been growing relative to the population for Kentucky as well as to the population of neighboring states. While state employment has been declining relative to population, local employment has been growing at a rate that more than replaces the declines in state government employment.

STATE AND LOCAL GOVERNMENT EXPENDITURES AND EMPLOYMENT BY GOVERNMENT FUNCTION

This section provides several alternative comparisons between Kentucky and its neighbors on costs and resources used in four different government functions: central administration, financial administration, primary and secondary education, and highways. As discussed previously, examining different government functions individually is important because states differ in both how they allocate expenditures across functions and between state and local governments. As a result, for some of the functions observed, our primary focus is on combined state and local expenditures rather than on state expenditures alone. To facilitate comparisons over time, we report inflation-adjusted amounts (2002 dollars) as in the preceding section. In addition to reporting per capita spending, we rank Kentucky relative to the other states and calculate the annualized change in real (inflation-adjusted) government spending on the function over our period of analysis.

Differences in per capita spending by government function or service are not, by themselves, indications of differences in efficiency or performance. These differences could be explained by differences in the costs of production of the services in the states, differences in use, and, possibly, differences in the quality or extent of the services provided. It is difficult to quantify, at least in a relatively simple and direct way, these differences for some services. However, for other services and functions, we can at least provide some indication of differences in the use of services, that is, some measure of output. Thus, for primary and secondary education, we report expenditures per student, and, for corrections, we report expenditures per inmate. For highways, we report expenditures per mile of highway. While these measures still do not account for differences in the quality or effectiveness of the government service or differences in costs of production, they are undoubtedly a better baseline than expenditures per capita.

We can also obtain insights into the production of government services by examining employment and compensation within the government function. As we calculated for expenditures, we determine employees per 1,000 residents for each function and, where possible, clients per employee. For example, for primary and secondary education, we calculate students per faculty member, and, for corrections, inmates per employee.

Central Administration

Expenditures on the central administration of state and local government are not related to the provision of any specific government function nor are they related to financial administration, as expenditures by the revenue function are. Instead, these expenditures are related to the general operations of the executive and legislative branches of government. For this reason, we make no attempt to measure an "output" or "quality of services" associated with central administration; instead we provide comparisons and trends based on per capita expenditures. When comparing central administration expenditures, particularly on a per capita basis, it is important to bear in mind that these services are likely to exhibit economies of scale. That is, while central administration costs can be expected to increase with the population of a state, they are not likely to increase at the same rate as the population.

Formally, the U.S. Census Bureau's *Annual Survey of Government Finances and Employment*, the source of our data, defines government administration, which we refer to as central administration, as "[g]overnment-wide executive,

Table 9

State and Local Expenditures on Central Administration, Selected Years

	Per capita, 2002$			Rank			
	1992	1997	2002	1992	1997	2002	Annual % change
Kentucky	35	49	69	8	5	4	7.2
United States	51	54	63	3	3	5	2.2
Illinois	58	49	90	2	6	1	4.5
Indiana	66	68	85	1	1	2	2.5
Missouri	36	52	56	6	4	6	4.5
Ohio	35	39	51	7	9	7	4.0
Tennessee	30	41	43	9	8	9	3.6
Virginia	41	47	51	5	7	8	2.3
West Virginia	42	65	80	4	2	3	6.6

Table 10

State and Local Employment in Central Administration Per 1,000 residents, Selected Years

	Per 1,000 residents			Rank			
	1992	1997	2002	1992	1997	2002	Annual % change
Kentucky	0.81	0.90	1.01	4	4	1	2.29
United States	0.86	0.94	0.94	3	2	4	0.92
Illinois	0.95	0.99	0.93	1	1	5	–0.27
Indiana	0.88	0.94	1.00	2	3	2	1.23
Missouri	0.61	0.83	0.84	8	7	7	3.23
Ohio	0.69	0.86	0.94	6	5	3	3.15
Tennessee	0.60	0.74	0.81	9	9	9	3.14
Virginia	0.76	0.84	0.83	5	6	8	0.93
West Virginia	0.66	0.81	0.88	7	8	6	2.85

administrative, and staff service agencies other than financial, judicial, legal, and Federal or state legislative activities."[4]

For example, costs associated with the legislative and executive branches of government are only weakly linked to population, as the number of legislators, support staff, and executive branch personnel are not likely to be significantly greater in larger states.

[4] For the definitions and examples from the manual for the *Annual Survey of Government Finances and Employment*, see www.census.gov/govs/www/classfunc29.html.

Table 9 gives the combined state and local central administration spending per capita. In 2002, Kentucky ranked fourth among the states, with spending of $69 per capita. Given the more centralized nature of Kentucky's government structure, the higher ranking for state spending is no surprise. Central administrative costs per capita are a small share of state and local government expenditures and, therefore, have a relatively modest influence on total state or combined spending. It is perhaps more important, in the case of Kentucky, to consider the rate at which central administrative expenditures have been

Table 11

State and Local Expenditures on Financial Administration, Per Capita, Selected Years

	Per capita, 2002$			Rank			
	1992	1997	2002	1992	1997	2002	Annual % change
Kentucky	73	70	80	7	8	7	0.9
United States	92	104	114	3	5	5	2.1
Illinois	81	122	115	5	4	4	3.6
Indiana	74	79	100	6	6	6	3.0
Missouri	65	71	76	8	7	8	1.6
Ohio	102	124	191	1	2	1	6.5
Tennessee	53	64	67	9	9	9	2.5
Virginia	98	124	119	2	1	3	1.9
West Virginia	92	123	174	4	3	2	6.6

Table 12

State and Local Employment in Financial Administration Per 1,000 Residents, Selected Years

	Per 1,000 residents			Rank			
	1992	1997	2002	1992	1997	2002	Annual % change
Kentucky	1.14	1.28	1.17	6	6	6	0.22
United States	1.25	1.36	1.33	4	5	4	0.60
Illinois	1.13	1.14	1.10	8	8	9	–0.27
Indiana	1.40	1.50	1.17	2	2	7	–1.83
Missouri	1.15	1.21	1.18	5	7	5	0.33
Ohio	0.99	1.44	1.45	9	4	3	3.87
Tennessee	1.13	1.14	1.15	7	9	8	0.21
Virginia	1.48	1.52	1.53	1	1	2	0.37
West Virginia	1.33	1.48	2.13	3	3	1	4.85

increasing. Per capita state and local spending in Kentucky increased during the 10-year period from 1992 to 2002 by an inflation-adjusted rate of 7.2 percent during this period, the highest rate among all the comparison states.

Table 10 reports state and local employment in central administration per 1,000 residents. Kentucky has the highest ranking in this category. While Kentucky ranks first in combined state and local employment, the differences between Kentucky and the rest of the states (and the U.S. average) in this category are not very pronounced.

Financial Administration

Financial administration includes government services provided by the finance and administrative agencies of government and revenue-collection and auditing/accounting agencies. As with central administration, output is difficult to measure for these services. Although it would seem reasonable to expect that states with smaller populations might spend more per capita, based on an expectation of economies of scale in these services, examination of costs for Kentucky and its neighbors does not seem to suggest that this is the case.

The definition of "financial administration" guiding the collection of data for the U.S. Census Bureau's *Annual Survey of Government Finances and Employment* is "[o]fficials and central staff agencies concerned with tax assessment and collection, accounting, auditing, budgeting, purchasing, custody of funds, and other finance activities."[5]

Table 11 shows that Kentucky spends relatively less in state and local expenditures on financial administration, ranking near the bottom of the comparison states and having a real per annum increase of only 0.9 percent. Not surprisingly, combined state and local employment for Kentucky, however, ranks low, with a rate of state and local financial employment of 1.17 per 1,000 residents, which is similar to most of its surrounding states (Table 12).

Primary and Secondary Education

While the provision of primary and secondary education is the responsibility of local governments, specifically school districts, it is heavily financed by state funds. In Kentucky, in 2001, 67 percent of primary and secondary education funding came from state sources, far above the typical level for its neighboring states with the exception of West Virginia. The state government is also involved in primary and secondary education through its regulatory role, imposing requirements for training, curricula, and facilities.

Although we use a rather standard measure of output for education (i.e., number of students), this measure, as with other measures of output we have used, does not adjust for the quality of services. In particular, higher expenditures per student may indicate a better quality education, a less efficient provision of services, or, possibly, both. Here, we make no attempt to measure the quality of services provided to students or to provide output measures such as results on standardized tests. While these issues are certainly important in understanding the efficacy of educational services, they are beyond the scope of this study.

Table 13 provides a comparison of primary and secondary education costs per student (average daily attendance) for Kentucky and its neighbors for 1992, 1997, and 2002. Current expenditures, including all expenditures except capital expenditures, are reported. Administration and instructional expenditures are reported separately. As the table shows, educational costs per student are quite low in Kentucky when compared with its neighboring states; Kentucky ranks seventh in both current expenditures and instructional expenditures per student in 2002. Administrative spending per student is relatively higher—in fact, the highest among the states in 1997, although the rank decreased to fifth in 2002.

Table 14 reports (i) student-to-teacher, (ii) student-to–administration and staff, (iii) student-to–central administration and staff, and (iv) student-to–central administration ratios for Kentucky and its neighboring states. For this table, bear in mind that a higher student-to-teacher or student-to-administrator ratio means fewer employees per output. Thus, the higher (closer to 1) the state ranks, the fewer the number of employees per student. As Panel A of the table shows, Kentucky has relatively high student-to-teacher ratios and there have been very modest decreases in the number of students per teacher during the period 1992 to 2002. In contrast, the ratio of students per administrators including staff (Panel B) was the second lowest among the states in 2002 and decreased at a rate of 5.2 percent per annum from 1992 to 2002. This is by far the greatest decrease in the ratios of students to administrators among Kentucky and its neighboring states. Panel C focuses on the ratio of students per district central office administrators and staff: While Kentucky has the third lowest ratio of students per central office administrators and staff, this ratio has decreased at a rate of 6.8 per annum since 1992. This represents the greatest increase in central administrators and staff (per student) among the states.

Focusing only on central administrators and not including staff (Panel D) shows more modest increases (in percentage terms) in central administrators, indicating the increase has been primarily staff and not administrators in central offices.

[5] The online version of the *Government Finance and Employment* manual has the definition of financial administration and examples at www.census.gov/govs/www/classfunc23.html.

Table 13

Current, Administrative, and Instructional Expenditures on Primary and Secondary Education, Various Years (2002$)

	Per student (average daily attendance)			Rank			
	1992	1997	2002	1992	1997	2002	Annual % change
A. Current expenditures							
Kentucky	6,051	6,646	7,536	7	5	7	2.22
Illinois	7,270	7,350	8,967	2	2	1	2.12
Indiana	6,506	7,403	8,268	4	1	4	2.43
Missouri	6,193	6,527	7,699	6	6	6	2.20
Ohio	7,301	7,305	8,928	1	4	2	2.03
Tennessee	4,734	5,617	6,489	8	8	8	3.20
Virginia	6,255	6,363	7,928	5	7	5	2.40
West Virginia	6,511	7,307	8,451	3	3	3	2.64
B. Administration							
Kentucky	604	629	648	3	1	5	0.71
Illinois	607	615	786	2	2	1	2.63
Indiana	481	548	621	6	6	6	2.59
Missouri	568	595	694	4	5	3	2.02
Ohio	618	612	779	1	3	2	2.34
Tennessee	371	417	459	8	8	8	2.16
Virginia	458	447	601	7	7	7	2.76
West Virginia	554	611	694	5	4	3	2.28
C. Instructional expenditures							
Kentucky	3,707	4,036	4,625	7	5	7	2.24
Illinois	4,355	4,421	5,335	1	3	1	2.05
Indiana	4,042	4,629	5,032	3	1	4	2.22
Missouri	3,757	4,006	4,690	5	6	6	2.24
Ohio	4,161	4,349	5,181	2	4	3	2.22
Tennessee	3,013	3,642	4,223	8	8	8	3.43
Virginia	3,725	3,865	4,887	6	7	5	2.75
West Virginia	3,939	4,526	5,212	4	2	2	2.84

SOURCE: National Center for Educational Statistics, U.S. Department of Education (http://nces.ed.gov/).

Table 14

Student-to-Teacher and Student-to-Administrator Ratios, Selected Years

A.	Student to teacher (full-time equivalent)			Rank			
	1992	1997	2002	1992	1997	2002	Annual % change
Kentucky	17.3	16.5	16.3	3	4	2	–0.6
Illinois	16.8	16.8	15.9	5	2	3	–0.5
Indiana	17.6	17.2	16.7	2	1	1	–0.5
Missouri	16.2	15.0	13.9	6	6	7	–1.5
Ohio	16.9	16.7	14.7	4	3	5	–1.4
Tennessee	19.6	16.5	15.8	1	4	4	–2.1
Virginia	15.1	14.3	11.8	8	8	8	–2.4
West Virginia	15.2	14.4	14	7	7	6	–0.8

B.	Student to administration and staff			Rank			
	1992	1997	2002	1992	1997	2002	Annual % change
Kentucky	87.3	94.6	51.1	4	1	7	–5.2
Illinois	99.3	85.1	74.9	1	4	4	–2.8
Indiana	93.3	90.3	89.5	2	3	1	–0.4
Missouri	64	58.4	71.1	7	8	5	1.1
Ohio	58.4	66.3	46.8	8	7	8	–2.2
Tennessee	77.9	71.7	70.7	6	6	6	–1.0
Virginia	91.4	92.2	77	3	2	2	–1.7
West Virginia	86.9	83.5	77	5	5	2	–1.2

C.	Student to district central administration and staff			Rank			
	1992	1997	2002	1992	1997	2002	Annual % change
Kentucky	206.5	184.6	102.1	5	5	6	–6.8
Illinois	270.8	210.8	184.3	3	4	4	–3.8
Indiana	727.8	662.8	637.4	1	1	1	–1.3
Missouri	170.7	106.1	93.8	6	8	8	–5.8
Ohio	121.4	122.1	94.4	8	7	7	–2.5
Tennessee	257.8	217.6	237.2	4	3	2	–0.8
Virginia	292.7	317.4	226.4	2	2	3	–2.5
West Virginia	152.1	133.4	127.3	7	6	5	–1.8

D.	Student to district central administration			Rank			
	1992	1997	2002	1992	1997	2002	Annual % change
Kentucky	646.6	635.6	543.4	6	4	6	–1.7
Illinois	1,128.7	572.1	517.3	2	6	7	–7.5
Indiana	1,086.7	1,072.6	1,031.7	3	1	1	–0.5
Missouri	1,020.6	831.6	701.4	4	3	3	–3.7
Ohio	322.8	333.4	280.7	8	8	8	–1.4
Tennessee	1,019.3	504.8	775.3	5	7	2	–2.7
Virginia	573.9	634.8	634.6	7	5	5	1.0
West Virginia	1,233.7	936.1	680.6	1	2	4	–5.8

Table 15

State and Local Expenditures on Highways, Selected Years

	Per capita, millions of 2002$			Rank			
	1992	1997	2002	1992	1997	2002	Annual % change
Kentucky	368	365	477	3	3	2	2.6
United States	340	349	402	6	6	6	1.7
Illinois	399	359	451	1	4	3	1.2
Indiana	273	311	330	9	9	8	1.9
Missouri	320	358	436	8	5	4	3.2
Ohio	322	315	359	7	8	7	1.1
Tennessee	341	338	306	5	7	9	–1.1
Virginia	359	420	426	4	2	5	1.7
West Virginia	392	513	576	2	1	1	3.9

Kentucky's ratio of students to school administrators was also the second lowest among states in 2002, and, during that period, the rate of reduction in that ratio, 5.2 percent per annum, was again the greatest among our benchmark states.

It is possible to calculate a salary figure for employees in primary and secondary education and even calculate a salary figure for personnel involved in instruction. However, we cannot calculate the salaries of specific educational occupations such as administrator or teacher because administrative staff are included in salary expenses for administrators and instructional aides are included in instructional salaries. For this reason, we do not attempt to construct any salary figure and, instead, focus on educational spending and employment as it relates to the number of students being taught.

Highways and Roadways

While revenues for highway and roadways in most states, including Kentucky, do not come from general funds, they are still a major expenditure for the state and a critical component of infrastructure. Spending in Kentucky, as is the case with most services, is primarily done by the state government, with state expenditures comprising 80 percent of combined state and local spending in 2000. In contrast, the U.S. average is only 60 percent; in Illinois it is less than 50 percent. Thus, when reporting state expenditures and employment, clearly, appropriate comparisons require comparisons of state and local expenditures and employment. Table 15 reports combined state and local governments capital outlays per capita. For both state and state and local per capita spending, Kentucky ranks second, trailing only West Virginia.

It is difficult and probably misleading to attempt to infer much about relative costs of or efficiency in the production of highway services based on per capita costs. Per capita costs could vary for a number of reasons that are unrelated to efficiency in provision, including differences in highway miles (per capita), terrain, climate, and usage. While all these factors are likely to influence costs, attempts to account for all of them are beyond the scope of this study. However, we do attempt to account for differences in highway usage and highway miles using data from the Highway Performance Monitoring System (HPMS) administered by the Federal Highway Administration (www.fhwa.dot.gov/policy/ohpi/hpms/). Table 16 reports usage (average annual daily traffic flow) and lane miles for each of the states for federal, state, and local highways and roadways. As the table indicates, Kentucky has significantly more lane miles, particularly controlled by the state, than many states with much larger populations.

Table 16

Average Daily Traffic Flow and Lane Miles by Government in Control, 1999

	Average daily traffic			Lane miles		
State	Federal	Local	State	Federal	Local	State
Kentucky	2.00	1.42	0.81	2,053	100,720	60,812
Illinois	9.79	2.35	2.16	511	244,485	43,952
Indiana		0.45	0.56		166,332	28,248
Missouri	1.12	0.49	0.63	2,208	181,739	69,938
Ohio		4.28	1.88	540	193,218	55,681
Tennessee		3.37	2.06	594	147,821	35,825
Virginia	2.02	7.21	3.21	3,793	26,335	122,929
West Virginia		4.21	1.41	1,355	4,528	70,233

SOURCE: Highway Performance Monitoring System (HPMS) administered by the Federal Highway Administration (http://www.fhwa.dot.gov/policy/ohpi/hpms/).

Table 17

Highway Expenditures, Per Mile Traffic Flow in 1999, 2002$

	Road miles			Lane miles		
State	Local	State	Combined	Local	State	Combined
Kentucky	7	55	24	4	25	12
Missouri	11	42	19	5	19	9
Indiana	9	121	23	5	48	11
Tennessee	10	90	23	5	36	11
West Virginia	24	27	26	12	13	13
Illinois	20	117	32	10	45	15
Ohio	21	102	36	10	41	17
Virginia	44	41	41	20	19	19

Table 17 reports the expenditures per traffic mile for state, local, and combined (state and local) highways for fiscal year 1999. This is calculated using the data in Table 16 with data on highway expenditures (Table 15). Traffic miles are simply the number of miles of roadways and the average annual traffic flow. Costs are reported both per mile of roadway and per mile of lanes. As the table suggests, once differences in use and miles of roadway are accounted for, Kentucky's costs are relatively low.

Table 18 reports highway employees per 1,000 residents. Again, when costs are measured in terms of population, Kentucky has high levels of employment. Although not reported here, if costs per mile of roadway is calculated, Kentucky has relatively modest employment in this function.

THE DETERMINANTS OF STATE AND LOCAL EXPENDITURES

In preceding sections, we documented the differences in state and state and local expenditures for Kentucky and its neighboring states.

Table 18

State and Local Employment in Highways per 1,000 Residents, Selected Years

	Per 1,000 residents			Rank			
	1992	1997	2002	1992	1997	2002	Annual % change
Kentucky	2.25	2.18	2.13	5	4	3	–0.56
United States	2.06	2.00	1.90	6	6	6	–0.81
Illinois	1.56	1.61	1.69	9	9	9	0.81
Indiana	1.85	1.81	1.69	8	8	8	–0.87
Missouri	2.26	2.52	2.34	4	2	2	0.33
Ohio	1.85	1.86	1.84	7	7	7	–0.07
Tennessee	2.27	2.19	1.92	3	3	5	–1.66
Virginia	2.37	2.08	1.98	2	5	4	–1.78
West Virginia	3.65	3.76	3.26	1	1	1	–1.11

However, we made no attempt to discern what might be the reason for these differences in expenditures and, in particular, whether these differences might be due to differences in the cost of provision or in the quality (and mix) of services provided. In particular, if there are differences in costs, are these differences in costs attributable to factors external to government agencies, the providers of services, or are they related to factors that might be considered internal to the operation and structure of the state and local governments?

Definitive answers to these questions are beyond the scope of this study, particularly given the lack of a measure of quality of services or, for some functions, even a measure of quantity or customer base. However, we believe that by estimating the relationship between expenditures and factors likely to affect it, we can offer some insights into understanding some of this variation in expenditures among states.

Data

We estimate the relationship for a few categories (including total state and local spending, administrative spending, and primary and secondary education) and what we believe are factors likely to influence spending, both because of supply (cost) and demand considerations. Data on these categories of government spending are obtained for the years 1992, 1997, and 2002 from the Census of Governments. All spending is converted to 2002 dollars and measured on a per capita basis, with state and local spending combined.

Table 19 shows variable means and values for the categories and influencing factors. Explanatory variables include the state population (and population squared) and population density (people per square mile). How the population affects per capita costs and, in particular, whether there is evidence of economies or diseconomies of scale related to the population of the state is a primary focus of this exercise. While our primary focus is on population rather than population density, we estimate a model in which both measures are included.

In addition, the distribution of expenditures between state and local governments is included with expenditure share measuring the fraction of expenditures in the category made by the state government. Our intention in including this variable is to see whether more centralized government service provision results in greater spending, perhaps because of reduced monitoring or less Tiebout competition. In addition to how the expenditures are distributed between state and local governments, we also have data on the number of counties, municipalities, and school districts and determine the average population for each of these jurisdictions for each state. This, we believe, offers some measure of whether there are economies or diseconomies of scale associated

Table 19

Variable Means and Values for Eight States

	Mean, eight states	Kentucky	Illinois	Indiana	Missouri	Ohio	Tennessee	Virginia	West Virginia
Total expenditures	6,372	6,073	6,945	5,895	5,816	7,009	6,325	6,006	6,609
Administrative expenditures	247	330	309	229	408	215	310	352	308
Population (1,000)	7,243	4,089	12,585	6,158	5,679	11,410	5,792	7,273	1,805
Density	165	103	226	172	82	279	141	184	75
Expenditures, state share	0.63	0.74	0.56	0.61	0.63	0.66	0.55	0.64	0.79
County population	77,556	34,370	123,384	67,674	49,823	129,664	62,960	76,564	32,822
Municipal population	13,543	9,646	9,748	10,861	6,004	12,113	16,597	31,762	7,715
Students per district	3,782	3,339	1,954	3,056	1,717	2,241	6,702	5,845	4,963
Federal revenue to state and local	0.21	0.24	0.17	0.18	0.24	0.19	0.23	0.16	0.26
Local revenue to state	0.01	0.00	0.02	0.01	0.01	0.01	0.01	0.01	0.01
State revenue to local	0.31	0.34	0.28	0.30	0.29	0.35	0.21	0.33	0.42
Relative earnings	0.91	0.82	1.08	0.88	0.89	0.92	0.87	1.00	0.76
Employment to population	0.59	0.57	0.59	0.59	0.62	0.60	0.60	0.62	0.49
Poverty rate	11.7	14.8	11.3	9.6	11.3	10.2	13.6	9.6	16.1
Income, median	40,789	35,875	44,946	41,973	40,309	42,246	37,129	48,224	30,695
Unemployment rate	4.9	6.1	5.9	5.1	5.0	4.7	4.8	4.5	4.6
Lower house, % Democrat	0.52	0.66	0.53	0.53	0.53	0.40	0.58	0.34	0.75
Upper house, % Democrat	0.48	0.47	0.46	0.36	0.41	0.33	0.55	0.45	0.82
Governor, Democrat	0.43	1	0	1	1	0	0	0	1
African American	0.12	0.07	0.15	0.08	0.11	0.11	0.16	0.19	0.03
Hispanic	0.04	0.01	0.12	0.04	0.02	0.02	0.02	0.05	0.01
Native American	0.002	0.002	0.001	0.002	0.004	0.002	0.002	0.003	0.002
Asian American	0.02	0.01	0.03	0.01	0.01	0.01	0.01	0.04	0.01
Urban	0.70	0.56	0.88	0.71	0.69	0.77	0.64	0.73	0.46
Median age	36.2	36.1	35.2	35.4	36.1	36.1	36.2	35.4	38.9

with the number and population of local governments. Differences in revenue sources are also considered, with variables included that measure the fraction of state and local revenue from federal sources, the fraction of state revenues from local sources, and the fraction of local revenues from state sources. These variables might reflect cost-sharing between the levels of government in the form of matching grants, for example, which might reduce the cost of providing the service for the government receiving the revenue. These data are also obtained from the Census of Governments.

Our next set of variables includes measures of employment and income, specifically, median household income, the (average annual) unemployment rate, the poverty rate, and a constructed variable, that is, the ratio of employment to population.[6] Additionally, in some specifications we also included measures of the political climate in the state, specifically the percentages of the lower and upper house members that were members of the Democratic Party and a categorical variable for the political party of the governor.[7] A final set of variables controlled for demographic factors, including the racial composition of the state's population (percentage of African American, Native American, and Asian American), the percentage of its population of Hispanic ethnicity, the percentage of the population living in urban areas, and the median age of the population.

For our estimates on the determinants of primary and secondary educational spending, our dependent variable is state and local education per student (in 2002 dollars) rather than per capita. Additionally, we include variables measuring student achievement: specifically, relative state scores on the National Assessment of Educational Progress (NAEP) exams for fourth graders in mathematics and reading and for eighth graders in mathematics. State averages for the SAT verbal and mathematics sections are also included, as is the percentage of students taking the exam in the state. Because states vary greatly in the percentage of students taking the exam and the exam is voluntary, this measure of achievement is probably less reliable than the NAEP, which is given to all students, with few exceptions, in participating states. Also included are measures of the educational attainment—that is, the percentage graduating from high school and the percentage having a BA or greater within the state population (adults over age 18), as this may affect the demand for educational services. We also include the percentage of the population between the ages of 5 and 19 and the percentage of primary and secondary students attending private schools, as these variables are likely to influence the (tax) cost of public education.

Empirical Methodology: Fixed-Effect Estimation

The empirical models we use are intended to address our two primary interests: to what extent are differences in costs related to economies or diseconomies of scale and how much of the difference in costs among states is not explained by differences in population or other factors that may influence costs. The basic form of the model we estimate is

$$E_{it} = \beta_1 P_{it} + \beta_2 D_{it} + \beta_3 S_{it} + \beta_4 J_{it} + \beta_5 R_{it} + \beta_6 IE_{it} + \beta_7 L_{it} + \beta_8 RH_{it} + \beta_9 T_t + \mu_i + \varepsilon_{it},$$

where the subscript i denotes the state and the subscript t denotes the year. The term E_{it} is the measure of expenditure per capita in state i and year t; P_{it} is state population (and population squared) in year t; D_{it} refers to population density; S_{it} is the state share of expenditures; J_{it} is a vector consisting of measures of population per jurisdiction (county, municipality, or district) or in some cases the number of jurisdictions; R_{it} measures sources of revenue; L_{it} are our measures of political sentiments; and RH_{it} is a vector of demographic variables reflecting race, ethnicity, age, and extent of urbanization in the state. The term T_t consists of year dummies. We estimate this as a fixed-effect model with μ_i being the state fixed-effect invariant over time by varying among states. In addition,

[6] Employment, unemployment, and poverty data are from the Bureau of Labor Statistics, U.S. Department of Labor, small area surveys. No survey was undertaken in 1992, so 1993 data were used. Income estimates are from the Census Bureau, U.S. Department of Commerce.

[7] Obtained from various editions of the *Statistical Abstract of the United States*, U.S. Printing Office.

Table 20

Estimation of Total State and Local Expenditures Per Capita

	A		B		C		D		E	
Variable	Coefficient	*T*	Coefficient	*t*	Coefficient	*t*	Coefficient	*T*	Coefficient	*t*
Population	–0.000106	0.000	–7.36E-05	–0.57	–0.000102	–0.83	–0.0003703	–3.14	–0.0003004	–2.39
Population^2	3.04E-12	0.000	2.70E-12	1.07	2.85E-12	1.16	5.78E-12	2.71	5.09E-12	2.35
Density			–3.201617	–1.24						
Density^2			0.0002	1.41						
State share of expenditures					2,833	2.54	3,507	2.46	3,737	2.58
County population							–0.0001991	–0.53	–0.0001367	–0.36
Municipal population							–0.0132	–3.15	–0.0212	–2.47
District students							0.0092212	0.46	0.0254	0.98
Federal to state and local revenue							1,521	0.96	1,506	0.89
Local to state revenue							2,483	0.69	4,226	1.16
State to local revenue							–2840	–3.48	–2420	–2.79
Relative earnings							5,255	5.17	4,637	4.57
Employment to population							1,422	0.68	1,197	0.55
Poverty rate							–79.3	–2.91	–75.5	–2.67
Median income							–0.0360532	–1.42	–0.0325	–1.13
Unemployment rate							36.1	1.33	56.6	2.1
Lower house, % Democrat							–991	–2.58		
Upper house, % Democrat							250	0.67		
Governor, Democrat							52.6	1.07		
African American									–35.9	–0.01
Hispanic									–2,545	–0.57
Native American									–386	–0.02
Asian American									–3,774	–1.58
Urban									482	0.37
Median age									–119	–1.44
1992	–1,270	69.59	–1,308	–17.67	–1,228	–17.06	–1,300	–7.75	–1,632	–5.38
1997	–920	57.35	–940	–15.61	–874	–14.75	–1,000	–8.3	–1,032	–6.83
F	149.37		145.45		113.65		152.98		140.92	
R^2	0.988		0.988		0.9848		0.9924		0.9924	
MSE	250.85		249.67		244.7		189.66		192.26	

Table 21

State and Local Expenditures Per Capita, State Fixed Effects for Alternative Specifications

State	Difference from mean	A	B	C	D	E
Kentucky	–221	–386	–478	–822	–850	–849
Illinois	466	771	830	810	1,147	991
Indiana	–389	–402	–345	–515	–754	–744
Missouri	–562	–609	–808	–714	–1,196	–1,117
Ohio	504	779	1,034	554	1,370	1,078
Tennessee	189	139	105	284	–17	71
Virginia	–161	–114	–59	–203	–238	–30
West Virginia	174	–178	–279	606	538	599

there is an error component varying both over time and among states, ε_{it}.

We estimate and report a number of alternative specifications, basically extending the sets of variables included as regressors. In addition to reporting the coefficients from the regression, we report the fixed effect for each of our eight states, our measure of the difference in cost not explained by the regressors that is invariant to the state over time. We also determine the relationship between expenditures per capita (or other base) and population for the alternative specifications. Finally, we decompose the variation in costs among our states to provide an indication of the determinants of differences in costs among the states.

Results

Total State and Local Expenditures. Table 20 reports the results of the fixed-effect estimation of per capita total state and local expenditures for a number of alternative specifications. Only when relative earnings are included as an explanatory variable (specifications D and E) do the population variables, both independently and evaluated jointly, become significant. The state's share of expenditures has a significant positive impact on total expenditures: for example, a 10 percent (0.10) increase in the state share increases total expenditures per capita from approximately \$270 to \$480, depending on the specification. Although the average population of counties or students in school districts had little impact on expenditures, the negative and significant coefficients on the municipal population variable provides some evidence of economies of scale in municipal services. Somewhat surprisingly, the state to local revenue variable is negative and significant, suggesting that a greater percentage of local funding coming from the state results in reduced total expenditures. The relative earnings variable appears to have a strong impact on total expenditures, with 10 percent higher earnings resulting in increased expenditures per capita of \$370 to \$525.

Table 21 reports the fixed effect for each of the eight states. The first column simply gives the actual difference in total expenditures per capita for each state without controlling for any factors that might affect expenditures. For Kentucky, for example, combined state and local expenditures are \$221 below the mean of the eight states. When controlling only for population, (A), Kentucky's expenditures fall to \$386. The large decrease in the fixed effect, –\$386 to –\$822, that occurs when the state share of expenditures variable (A to C) is included suggests that this is an important element in explaining state and local expenditures in Kentucky.

In Table 22 we use specification E reported in Table 20 to decompose the differences in expenditures between the states. Thus, for example, the \$707 reported under population for Kentucky equals

Table 22

Sources of Differences Among States in State and Local Expenditures

Decomposition of E	Population	Expenditure share	No. of local governments	Source of revenue	Relative earnings	Income/ employment	Demographics
Kentucky	707	399	77	–55	–450	–33	43
Illinois	–1,435	–253	28	92	764	–49	–75
Indiana	251	–79	40	–17	–157	132	137
Missouri	396	–12	111	95	–111	84	76
Ohio	–1,161	82	–16	–125	37	60	118
Tennessee	362	–307	11	283	–219	–15	38
Virginia	–75	27	–333	–120	387	–75	13
West Virginia	1,663	542	160	–208	–701	–139	–306

$$-0.0030404\left(Population_{KY}-\overline{Population}\right)$$
$$+\left(5.09E-12\right)\left(Population_{Ky}^{2}-\overline{Population}^{2}\right)$$

using the coefficients from specification E. Then, for Kentucky, population and the state's share of expenditure act to increase costs while its lower relative earnings reduce costs. Conversely, for Illinois, its large population and decentralized expenditures reduce expenditures, but its higher relative earnings and its demographic composition increase expenditures.

Figure 1 illustrates the relationship between population and per capita expenditures for our alternative specifications. For our most parsimonious specifications, the impact of population is relatively small and statistically insignificant. However, when controlling for the impacts of relative earnings and other factors, there are pronounced economies of scale of a large magnitude.

Administrative Costs. We use the same methodology and the same variables to examine the determinants of combined state and local administrative costs per capita. We broadly define administrative costs to include the categories of financial administration, judicial, other administration (central and legislative), and building operations. Although this is a relatively small share of total state and local government expenditures, we are interested in these expenditures because of frequent suggestions of excesses or waste in administration.

The results of the fixed-effects estimation, reported in Table 23, are generally qualitatively similar to those for total expenditures—with some notable exceptions. First, both the population and population^2 variables are statistically significant for all specifications. Similar to what was found with total expenditures, increases in the municipal population variable reduced administrative costs, but the share of administrative costs between the state and local governments had no impact on the level of administrative costs. Surprisingly, the relative earnings variable was positive but statistically insignificant. Increases in both the poverty rate and median income reduced administrative expenditures, as did the fraction of the population that was Hispanic or Asian American.

Table 24 reports the fixed effects for the eight states for the alternative specifications reported in Table 23, and Table 25 reports the decomposition of the differences in spending among the states and is analogous to what was presented in Table 22 for total expenditures. Viewing both tables suggests that differences in the population explain much of the differences in administrative costs among states. For Kentucky, as seen in Table 25, based on specification E, the difference between Kentucky's population and the mean for the eight states leads to an estimated increase in administrative expenditures of $142, while demographic factors in the state reduce predicted costs by $113. Other factors have a much smaller influence on costs.

Table 23

State and Local Administrative Costs

	A		B		C		D		E	
Variable	Coefficient	T	Coefficient	t	Coefficient	t	Coefficient	T	Coefficient	t
Population	–0.00004	–2.73	–3.5E-05	–2.54	–3.55E-05	–2.61	–0.0000403	–2.78	–0.0000594	–4.19
Population^2	6.78E-13	2.55	6.64E-13	2.44	6.57E-13	2.43	7.32E-13	2.76	9.68E-13	3.90
Density			–0.054976	–0.2						
Density^2			1.92E-06	0.12						
Expenditures, state share					130.2822	1.06	119.519	0.68	150.6181	0.92
Counties, population per							–0.0000829	–1.80	–0.0000558	–1.33
Municipalities, population per							–0.0012165	–2.66	–0.001953	–2.87
Revenue, federal to state and local							–750	–3.8	–608	-3.17
Revenue, local to state							–153	–0.34	–149	–0.36
Revenue, state to local							29.3	0.29	49.3	0.51
Salary, relative to other states							163	1.30	34.8	0.30
Employment to population							342	1.31	429	1.73
Poverty rate							–3.34	–1.00	–6.41	–2.01
Income, median							–0.00549	–1.74	–0.0054779	–1.66
Unemployment Rate							3.28	0.97	3.56	1.15
Lower house, % Democrat							–19.6	–0.41		
Upper house, % Democrat							–22.4	–0.49		
Governor, Democrat							4.31	0.71		
African American									998	1.53
Hispanic									1,516	2.99
Native American									1,554	0.71
Asian American									–463	–2.04
Urban									227	1.52
Age, median									–0.6726	–0.07
1992	–83.5	–11.3	–83.8	–10.49	–81.0	–10.22	–123	–5.89	–77.1	–2.22
1997	–45.7	–7.48	–46.2	–7.11	–43.9	–6.72	–86.8	–5.83	–62.0	–3.59
F	53.45		51.58		44.64		47.56		52.3	
R^2	0.9653		0.9678		0.9621		0.9755		0.9791	
MSE	27.44		26.953		26.942		23.652		22.054	

Table 24

State and Local Administrative Expenditures Per Capita, State Fixed Effects for Alternative Specifications

State	Difference from mean	A	B	C	D
Kentucky	–43	–127	–127	–139	–70
Illinois	29	153	153	158	88
Indiana	–3	–28	–27	–28	–17
Missouri	–31	–70	–72	–68	–87
Ohio	72	178	181	171	255
Tennessee	–60	–99	–99	–86	–102
Virginia	30	30	31	32	–29
West Virginia	6	–38	–40	–40	–36

Analogous to Figure 1, Figure 2 provides a relationship between per capita administrative costs and state population. In this case, decreasing costs are exhibited throughout the range of population for the states and there is much less variation in the extent of these economies of scale among the alternative specifications.

Primary and Secondary Education. Unlike many of the other functions of state and local governments, there is a voluminous literature examining educational finance and educational productions; much of this literature has focused directly on the relationship between educational expenditures and educational "outputs," most frequently performance on standardized tests, but occasionally on other measures such as high school completion or earnings.

Given the extensive research on educational finance and returns to education, our contribution to this literature is minor; perhaps, though, it is valuable as an examination of the impacts of differences in the structure and financing of education among states. We follow the same general methodology as used for our examination of total and administrative expenditures, albeit using some different measures of scale economies and some measures of output (test results). In addition, we also include an alternative measure of the age distribution, the fraction of the population under age 19, because this will affect the (tax) cost of educational services. Also, because primary and secondary education is almost exclusively provided by local governments, we do not include any measure of the state share in educational expenditures.[8]

We are again interested in the issue of economies of scale, but, rather than relate costs to the states total population, we consider costs relative to students per district and students per school. Rather than measure per capita expenditures, we use educational expenditures per student as our dependent variable. The results of our fixed-effect estimation are found in Table 26. From the table we can see that there are generally significant effects of the number of students per district. The marginal effect,[9] evaluated at the mean of 8,249 students per district, is statistically significant at a level of 0.05 in all specifications. The marginal effect of students per school is significant at a level of 0.05 in specifications B and C but not in the other specifications. These results can be seen more clearly in Figures 3 and 4. In Figure 3, the relationship between the average number of students per school in the state and expenditures per student is clearly U-shaped with, for most specifications, the minimum approximately between

8 The notable exceptions being Hawaii, which has a single school jurisdiction and, of course, the District of Columbia. Also excluded are the political variables. These variables were included in some unreported estimation and were found to be insignificant and having little impact on the other variables' coefficients.

9 Specifically the test is whether $\beta_{SD} + 2 * \overline{SD} \beta_{SD^2} = 0$, where SD refers to students per district.

Table 25

Sources of Differences in Administrative Expenditures Among States

State	Deviation from mean	Population	Expenditure share	No. of local governments	Source of revenue	Relative salary	Income/ employment	Demographics	Explained	Difference between explained and actual
Kentucky	–60	142	16	10	–19	–3	3	–113	36	96
Illinois	22	–146	–9	5	16	6	–16	188	44	22
Indiana	2	25	–3	6	13	–1	9	–38	11	10
Missouri	–78	40	0	16	–21	–1	19	–32	21	99
Ohio	100	–120	3	0	11	0	4	–19	–120	–220
Tennessee	–92	36	–10	–5	–18	–2	14	5	20	112
Virginia	2	–9	1	–36	28	3	–16	84	55	53
West Virginia	45	175	18	14	–29	–5	–15	–189	–32	–76
Mean	308									

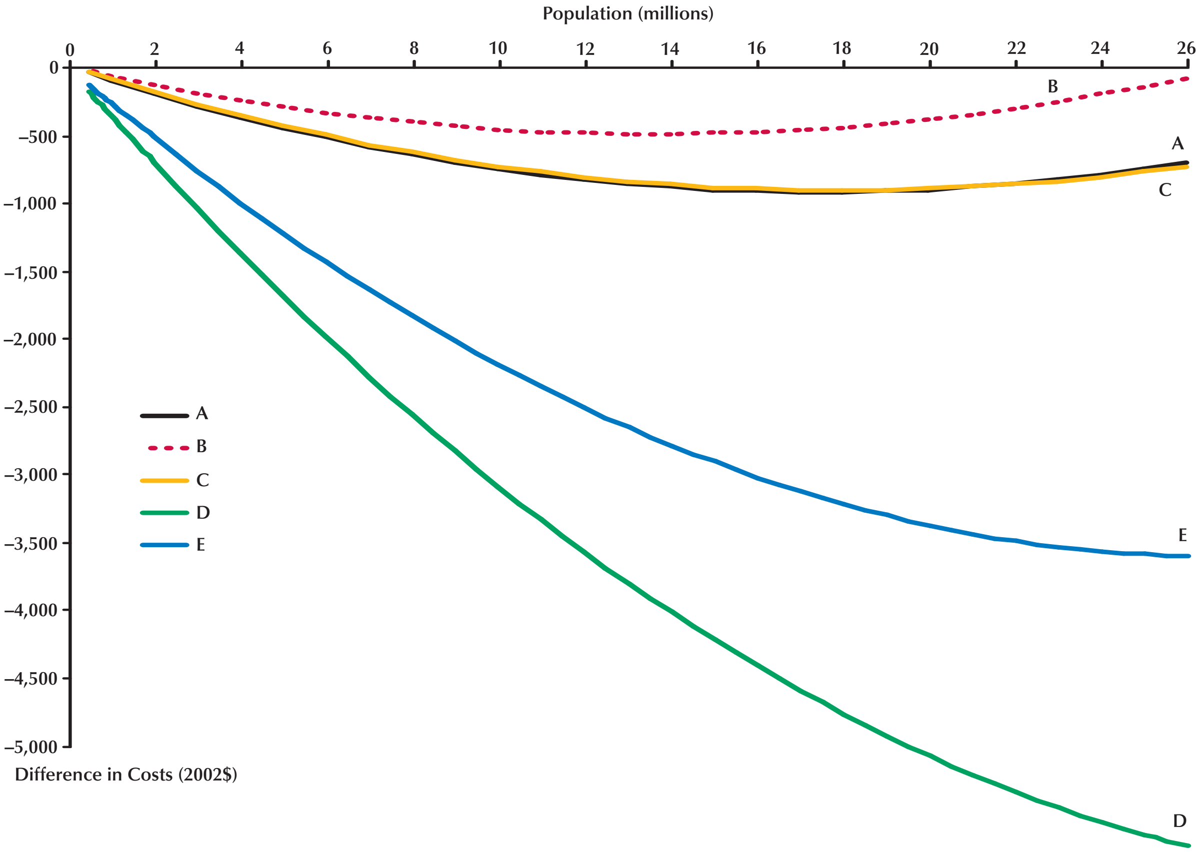
Table 26

State and Local Primary and Secondary Education Expenditures Per Capita, Fixed-Effect Estimation

	A		B		C		D		E	
Variable	Coefficient	T	Coefficient	t	Coefficient	t	Coefficient	T	Coefficient	t
Students/district	–0.0378313	–2.68	–0.1512172	–3.69	–0.1525052	–3.78	–0.1123068	–2.03	–0.1277763	–2.74
$(\text{Students/district})^2$	–2.24E-08	–0.16	2.16E-07	1.41	2.44E-07	1.6	1.89E-07	0.93	2.66E-07	1.52
Students/school	–11.28282	–1.6	–14.10065	–2.37	–12.4991	–2.11	–11.3965	–1.48	–7.901093	–1.21
$(\text{Students/school})^2$	0.0126233	2	0.0152521	2.86	0.013537	2.53	0.0122561	1.75	0.0100554	1.69
State to local revenue			2,121.976	2.31	1,780.62	1.93	1,433.601	1.07	1,147.321	1.17
Relative Earnings					2,136.813	1.93	3,530.235	1.64	3,185.971	2.27
Fraction < 19							–14,614.94	–1.45	–8,298.532	–1.03
NAEP, 4th reading							113.793	0.03		
NAEP, 4th math							5,016.547	0.97		
NAEP, 8th math							–6634.675	–1.03		
African American							3,108.104	0.3	6,643.901	0.72
Hispanic							–6,038.257	–0.83	–2,760.4	–0.44
Native American							35,067.22	0.96	29,748.77	1.03
Asian American							1,581.305	0.62	326.9808	0.14
Urban							207.0132	0.09	5.419015	0
Poverty rate							–27.48045	–0.56	–70.47135	–1.68
Median income							–0.0010409	–0.02	0.0097922	0.24
1992	–1,673	–22.02	–1,703.032	–23.8	–1,749.257	–23.47	–1,860.216	–6.76	–1,619.869	–6.53
1997	–1,368	–16.91	–1,301.528	–17.9	–1,343.076	–17.94	–1,474.479	–8.69	–1,398.308	–9.79
Number of observations	153		151		151		127		151	
F (56, 96)	60.33		74.06		74.96		57.73		67.98	
Probability > F	0		0		0		0		0	
R^2	0.9724		0.9784		0.9793		0.9859		0.9816	
Adjusted R^2	0.9563		0.9652		0.9662		0.9688		0.9672	
Root MSE	381.34		321.28		316.7		307.98		312.15	

Figure 1

Economies of Scale for Total State and Local Expenditures, Alternative Specifications



400 and 500 students, a range encompassing the mean for our sample, 454 students per school. In contrast, as shown in Figure 4, expenditures per student decrease throughout the relevant range for average students per district, with the difference in expenditures being quite substantial.

Both the magnitude and statistical significance of the coefficient on the state to local revenue variable were sensitive to specification, that is, inclusion of additional variables, with the magnitude of the coefficient decreasing with inclusion of demographic factors. This may not be surprising, as the extent of revenue sharing may be positively related to the degree of poverty and its geographical concentration within the state. Not surprisingly, the relative earnings variable has a positive and significant impact on expenditures, whereas the fraction < 19 variable has a negative impact but is only statistically significant in specification C. The relative average NAEP scores were insignificant, as were the demographic variables.

Table 27 reports the estimated fixed effect for each of the eight states. Although the magnitude of the fixed effect generally diminished with increased explanatory variables, as expected, West Virginia's effect increased dramatically when demographic characteristics were included and NAEP scores omitted. In contrast, Tennessee's fixed effect was relatively invariant to specification while Kentucky's decreased dramatically, from a value of –1,076 with fixed effects to only –365 in specification E. In contrast, Illinois's large positive

Figure 2

Economies of Scale for State and Local Administrative Expenditures, Different Controls

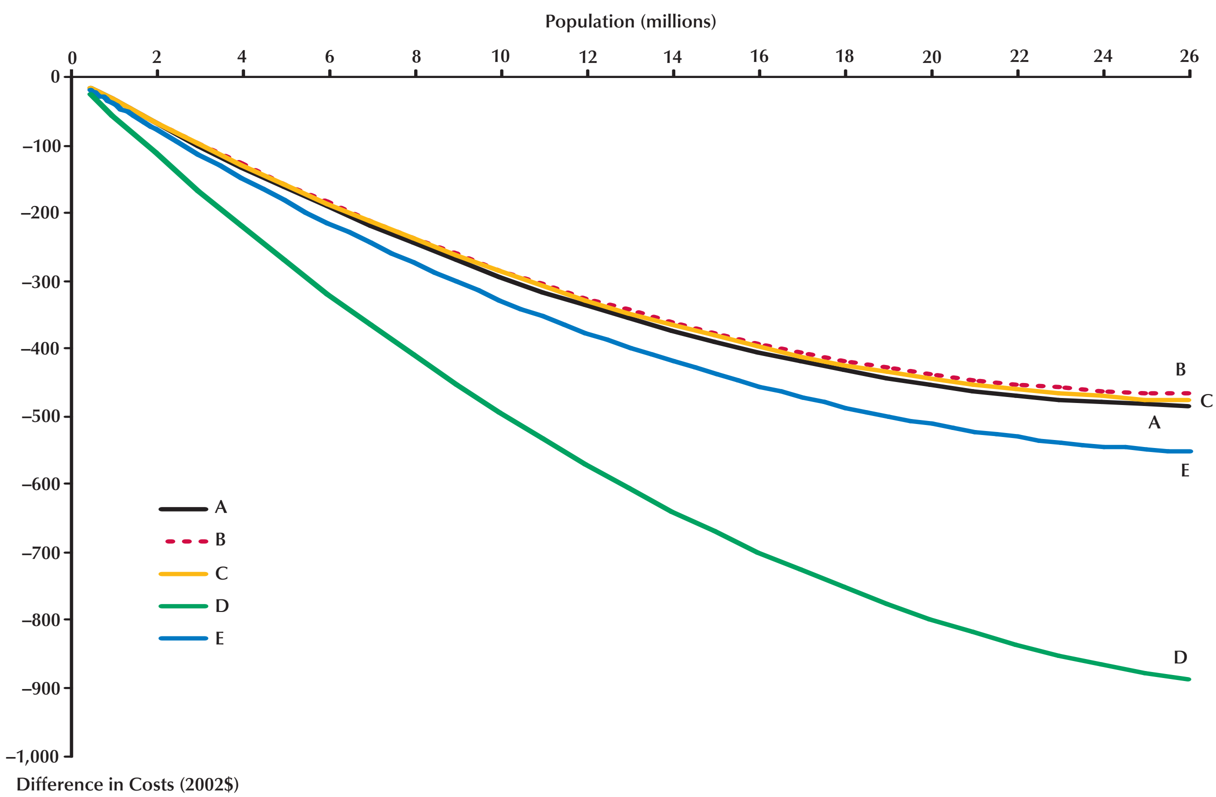


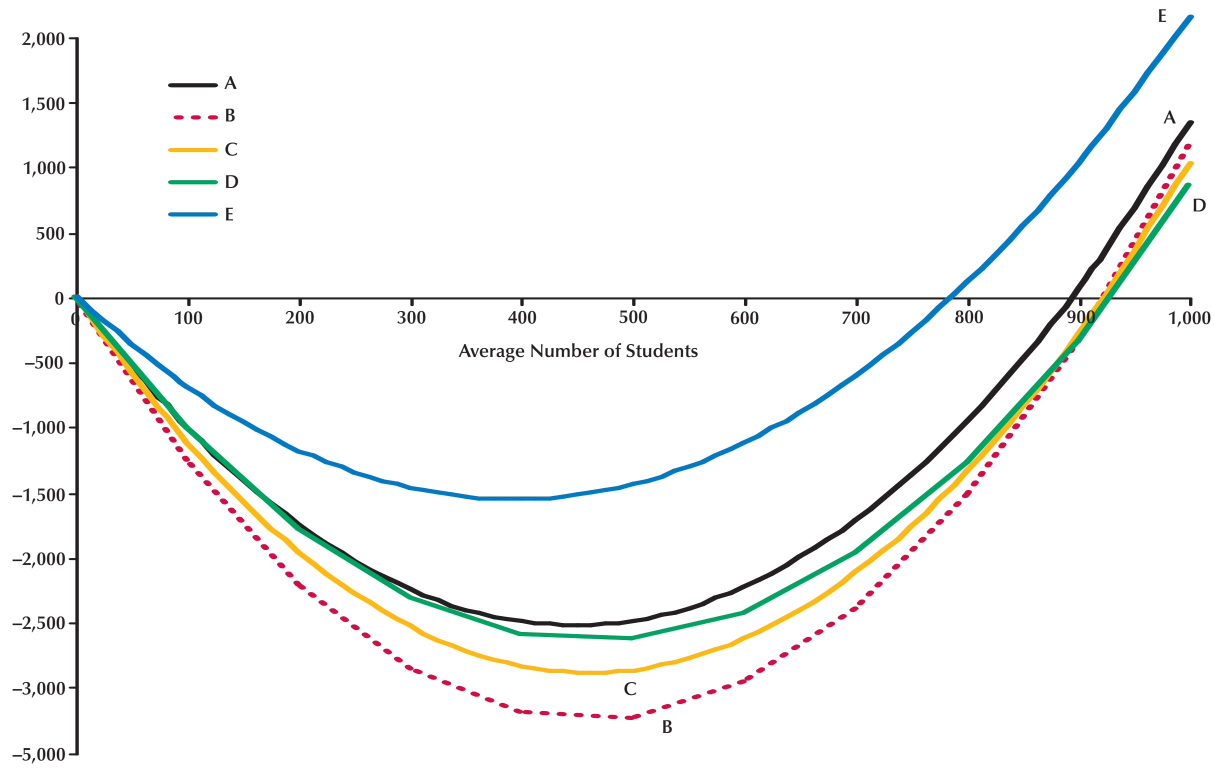
Table 27

Primary and Secondary Expenditures Per Capita, State Fixed Effects for Alternative Specifications

	Difference from mean	A	B	C	D	E	F
Kentucky	–1,076	–1,037	–1,225	–1,041	–989	–748	–365
Illinois	729	729	637	272	918	907	9
Indiana	415	422	353	381	342	386	392
Missouri	–204	–254	–461	–449	–214	–423	–324
Ohio	844	858	680	614	866	715	483
Tennessee	–1,412	–1,407	–863	–807	–1,057	–1,234	–1,268
Virginia	358	351	685	564	168	–93	–520
West Virginia	346	338	195	465	–33	490	1,594

Figure 3

Costs versus Average Number of Students Per School



effect became statistically insignificant in specification E.

Table 28 reports the results of a decomposition of specification E, explaining the difference in each state's predicted expenditures and the mean expenditure of the states. Illinois, Virginia, and West Virginia have significant reductions in expenditures attributable to the significant number of students per district, with the relatively small number of students per district in Indiana, Ohio, and Tennessee increasing their expenditures. Only for Illinois did the number of students per school have much impact on expenditures. Relative earnings had a large impact on expenditures in Indiana and West Virginia, and demographics played a significant role in increasing costs in Illinois and West Virginia and reducing costs in Indiana.

CONCLUSION

In our sample of Kentucky and its seven neighboring states, states with smaller populations and more centralized spending have higher per capita expenditure. State government employment tends to mirror expenditure, that is, less populous and more centralized states have greater employment per capita; but this pattern breaks down when state and local government employment is considered. For state government salaries, lower-wage states tend to have relatively high wages, but for state and local salaries there is no clear pattern relating government salaries and relative wages.

Regarding particular government functions, no clear patterns emerge for central administration expenditure and employment, though lower-wage states, especially Kentucky and West Virginia, tend to have high central administration salaries.

Figure 4

Costs versus Average Number of Students Per District

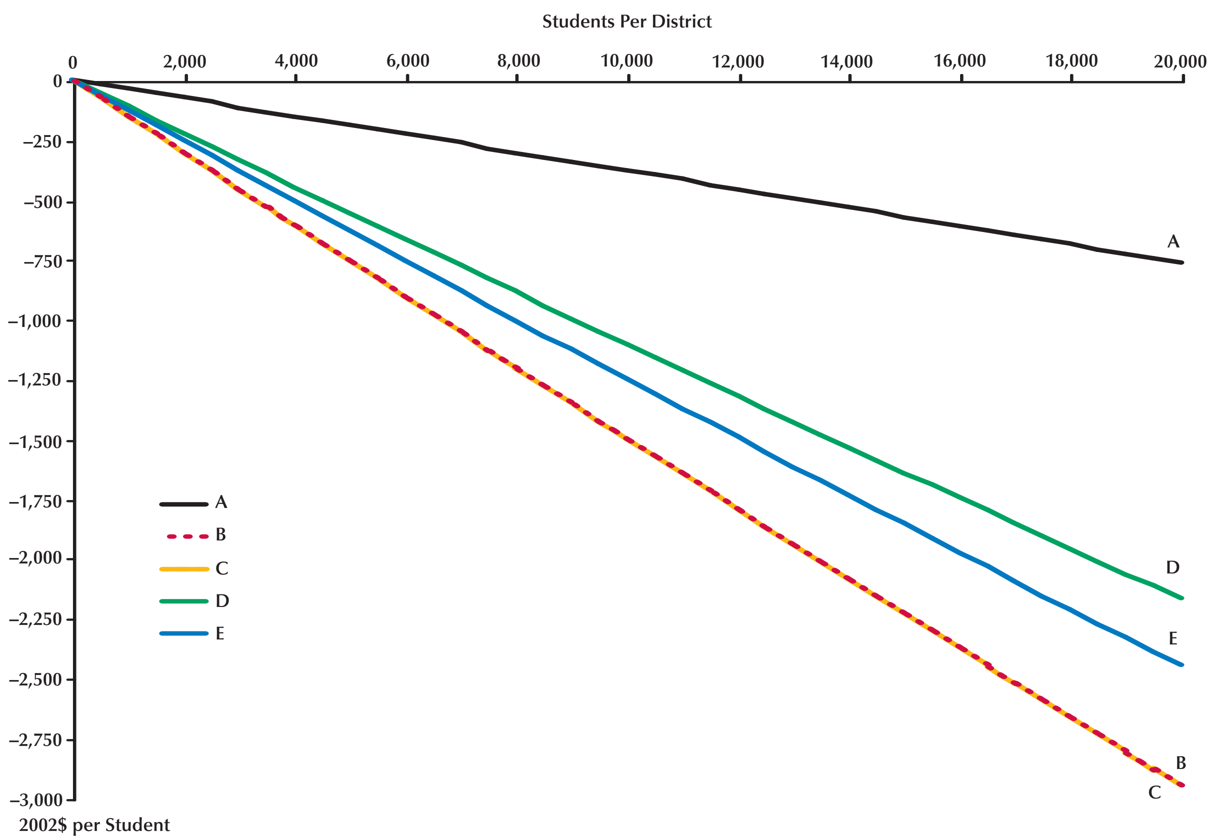


Table 28

Sources of Differences among States in Primary and Secondary Educational Expenditures

State	Districts	Schools	Revenue source	Relative salary	Age < 19	NAEP	Demographics
Kentucky	74	–45	–22	–77	–146	–13	–35
Illinois	–449	203	–7	122	–161	15	285
Indiana	197	–61	–54	624	–203	–26	–364
Missouri	43	–60	37	–300	39	–38	–126
Ohio	223	21	–40	–41	–125	–28	160
Tennessee	165	–63	48	71	–69	–36	146
Virginia	–329	52	–146	–124	61	81	182
West Virginia	–235	40	17	337	91	–14	297

Regarding primary and secondary education spending, with the exception of West Virginia, spending per student is higher for higher income, more populous states. All states experienced a reduction in the student-to-teacher ratio but also a reduction in the student-to-administrator ratio (with the exception of Missouri). The least populous states—West Virginia and Kentucky—have the highest per capita spending on highways. This does not hold for spending per road mile, however. West Virginia stands out as exceptionally high in employment in highway provision.

Our multivariate analysis reveals some interesting findings, too. There are economies of scale: More populous states have less spending per capita. States with more centralized spending have greater state and local total spending and higher-wage states have greater spending. States with a greater population per municipality and a higher poverty rate have lower spending. Controlling for more covariates tends to raise the estimated scale economy. The state fixed effects change substantially after controlling for the covariates. More populous states now tend to have higher expenditure. The results of the multivariate analysis for central administration spending tend to mirror the findings for total expenditure.

Regarding school expenditure: Economies of scale are strong for students per district, but less so for students per school. Higher-wage states have higher expenditure per student. Measures of student performance (i.e., test scores) have little relationship to spending. Control for covariates alters the estimated differences between states, but the ranking does not change much.

We find substantial differences in state spending, both in the aggregate as well as for specific functions. We also find substantial economies of scale in the provision of government services. Controlling for these economies of scale alters the ranking of high-to-low spending states. A major shortcoming of the study is that we have not controlled for or quantified the differences in the quality of public services among these states. However, by controlling for a number of factors that are likely to affect both the demand for and the cost of government services, we have reduced the difference in costs among states that is "unexplained," that is, differences in costs that cannot be attributed to differences in the demographics or populations of the states. It is the remaining "unexplained" difference in costs or residual that government officials who seek to claim efficiency must justify as representative of quality.

Forecasting Employment Growth in Missouri with Many Potentially Relevant Predictors: An Analysis of Forecast Combining Methods

David E. Rapach and Jack K. Strauss

In this paper, the authors examine different approaches to forecasting monthly Missouri employment growth in the presence of many potentially relevant predictors, including both regional and national economic variables. Following Stock and Watson (2003, 2004), they first generate simulated out-of-sample forecasts of Missouri employment growth at horizons of 3, 6, 12, and 24 months using individual autoregressive distributed lag (ARDL) models based on 22 potential predictors. They then consider 20 different methods from the extant literature for combining the forecasts generated by the individual ARDL models. At longer horizons of 12 and 24 months, combining methods based on Bayesian shrinkage techniques produce out-of-sample forecasts that are substantially more accurate than forecasts from an autoregressive (AR) benchmark model. Combining methods based on Bayesian shrinkage techniques also outperform simple combining methods (such as those that use the mean or median of the individual forecasts) at longer horizons. Nevertheless, simple combining methods consistently outperform the AR benchmark model at all horizons and appear to offer a low-cost way of generating reliable combination forecasts.

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1 INTRODUCTION

Bates and Granger's (1969) seminal work showed that combinations of individual forecasts often outperform individual forecasts. Well over a quarter century later, Stock and Watson (2004) analyze more than a dozen different methods for combining forecasts of output growth in the G7 countries and find that combination forecasts are a useful way of incorporating information from a large number of potentially relevant predictors. They show that combination forecasts of output growth often outperform forecasts generated by a benchmark autoregressive (AR) model and that simple methods, such as simple averaging or trimmed averaging of a large number of individual autoregressive distributed (ARDL) model forecasts, typically outperform more complicated methods. Combination forecast methods can exploit the information in a large number of potential predictors. This is especially relevant when forecasting a variable like national output growth, subject to both supply and demand shocks and possible instabilities in the data. In this case, it is difficult to know a priori which particular variables are the most relevant and, moreover, it is unlikely that a forecaster can specify a single econometric model that closely corresponds to the actual—and perhaps unknowable—data-generating process.

Although most of the literature, including

David E. Rapach is an assistant professor of economics and Jack K. Strauss is a professor of economics in the Department of Economics at Saint Louis University. The authors thank Michael Owyang, Jeremy Piger, Howard Wall, and BERG seminar participants at the First Annual Conference of the Business and Economics Research Group (BERG) of the Federal Reserve Bank of St. Louis for helpful comments. The authors acknowledge research support from the Simon Center for Regional Forecasting at Saint Louis University. The views expressed are the authors and do not necessarily represent the official positions of Saint Louis University.

Stock and Watson (1999, 2003, 2004), examines combination forecasts of national economic variables such as output growth and inflation, we are interested in the usefulness of combining methods when forecasting a regional economic variable. As discussed above, combining methods are likely to be useful when predicting aggregate economic variables because they incorporate information from a large number of potential predictors. This is also likely to be the case when forecasting regional economic variables, as a large number of both national and regional variables may contain information useful for forecasting. In the present paper, we consider forecasting employment growth in Missouri in the presence of a large number of potentially relevant predictors. Generating accurate forecasts of regional variables is important for planning purposes for businesses and state and local governments. Evaluating forecast combining methods for predicting a regional economic variable represents a natural complement to the extant literature on national economic variables.

We analyze forecasts of Missouri employment growth over the 1995:01–2005:01 out-of-sample period. This period includes the expansion of the late 1990s, the 2001 recession, and the subsequent "jobless" recovery, so it should represent an informative laboratory for analyzing forecasts of Missouri employment growth. We consider 22 potential predictors of Missouri employment growth, including 9 regional and 13 national economic variables. Following Stock and Watson (2003, 2004), we first generate simulated out-of-sample forecasts of Missouri employment growth from individual ARDL models, with each ARDL model based on 1 of the 22 potential predictors. We then use 20 different methods from the extant literature to construct combination forecasts of the individual ARDL model forecasts. The combining methods are based on the following: simple averaging using the mean, median, or trimmed mean (Stock and Watson, 2003, 2004); ordinary least squares (OLS; Granger and Ramanathan, 1984); weighted least squares (WLS; Diebold and Pauly, 1987); discount mean squared forecast error (MSFE; Stock and Watson, 2004); Bayesian shrinkage techniques (Clemen and Winkler, 1986; Diebold and Pauly, 1990); clusters formed on the basis of MSFE (Aiolfi and Timmermann, 2005); model selection (Swanson and Zeng, 2001); principal components (Chan, Stock, and Watson, 1999; Stock and Watson, 2004); approximate Bayesian model averaging (Draper, 1995); and exponential reweighting (Yang, 2004).

Previewing our results, we find that forecast combining methods can improve the forecasting of employment growth in Missouri, especially at longer horizons of 1 and 2 years. In particular, combining methods based on Bayesian shrinkage techniques generate forecasts that are up to 29 percent and 49 percent more accurate, in terms of MSFE, than the forecasts produced by an AR benchmark model at horizons of 1 and 2 years, respectively. Combination forecasts based on Bayesian shrinkage techniques also have an MSFE that is close to or below that of the best individual ARDL model forecast at horizons of 1 and 2 years, and Bayesian shrinkage combination forecasts have a lower MSFE than simple combining methods at these horizons. It should be noted that a number of the forecast combining methods fail to outperform the AR benchmark model, implying that it is critical to carefully select combining methods when forecasting Missouri employment growth. Simple combining methods appear to offer a low-cost way of generating reliable combination forecasts, as they consistently outperform the AR benchmark model at all horizons, in agreement with the findings of Stock and Watson (2004).

The rest of the paper is organized as follows: Section 2 describes the econometric methodology, Section 3 reports the empirical results, and Section 4 concludes.

2 ECONOMETRIC METHODOLOGY

2.1 Individual Forecasts

Let $\Delta y_t = y_t - y_{t-1}$, where y_t is the log-level of Missouri employment at time t, and let

$$y_{t+h}^{h} = (1/h)\sum_{j=1}^{h} \Delta y_{t+j} ,$$

so that y_{t+h}^{h} is the growth rate of Missouri employment over the next h months expressed at a monthly rate. Consider the following ARDL model:

$$(1) \qquad y^h_{t+h} = \alpha + \sum_{j=0}^{q_1-1} \beta_j \Delta y_{t-j} + \sum_{j=0}^{q_2-1} \gamma_j x_{i,t-j} + \varepsilon^h_{t+h},$$

where $x_{i,t}$ is one of the potentially relevant predictors ($i = 1,\ldots,n$), h is the forecast horizon, and ε^h_{t+h} is an error term. We consider forecast horizons of 3, 6, 12, and 24 months ($h = 3,6,12,24$). In order to form recursive simulated out-of-sample forecasts of y^h_{t+h} using equation (1), we first divide the sample into in-sample and out-of-sample portions, where the first R observations comprise the in-sample period and the last P observations make up the out-of-sample period. We compute the initial out-of-sample forecast for y^h_{R+h} based on the predictor $x_{i,t}$ as

$$\hat{y}^h_{i,R+h|R} = \hat{\alpha}_R + \sum_{j=0}^{q_1-1} \hat{\beta}_{j,R} \Delta y_{R-j} + \sum_{j=0}^{q_2-1} \hat{\gamma}_{j,R} x_{i,R-j},$$

where $\hat{\alpha}_R$, $\hat{\beta}_{j,R}$, and $\hat{\gamma}_{j,R}$ are the OLS estimates of α, β_j, and γ_j, respectively, in equation (1) using data through period R. We select the lag lengths (q_1 and q_2) in equation (2) using the Akaike information criterion (AIC) and data through period R considering a minimum lag length of 0 for q_1 and 1 for q_2 (thus ensuring that the potential predictor $x_{i,t}$ appears in equation (1)) and a maximum lag length of 12 for q_1 and q_2. We form the second out-of-sample forecast by updating the above process using data through period R+1. Continuing in this manner, we end up with a series of $P - (h - 1)$ simulated out-of-sample forecasts corresponding to the predictor $x_{i,t}$, $\{y^h_{i,t+h|t}\}^{T-h}_{t=R}$. Note that q_1 and q_2 are selected anew when computing each recursive out-of-sample forecast, so that the ARDL lag lengths in the forecasting equations can vary over time. We consider 22 potential predictors that define the individual ARDL models ($n = 22$). Apart from data availability and revisions, these simulated out-of-sample forecasts mimic the situation of a forecaster in real time.[1]

An AR model, equation (1) with the restriction $\gamma_j = 0$ for all j imposed, serves as the benchmark model. This is a common benchmark model when forecasting time-series variables. The AR model forecasts are computed recursively in a manner similar to the ARDL model forecasts, with the lag length selected by the AIC given a minimum (maximum) lag length of 0 (12). This produces a series of $P - (h - 1)$ simulated out-of-sample forecasts corresponding to the AR benchmark model, $\{y^h_{AR,t+h|t}\}^{T-h}_{t=R}$.

2.2 Forecast Combining Methods

We consider 20 different methods for combining the individual forecasts generated by the $n = 22$ ARDL models, and the methods can be organized into 10 different classes. Most of the forecast combining methods require a holdout period to calculate the weights used to combine the individual ARDL model forecasts, and we use the first P_0 out-of-sample forecast observations as holdout observations. All of the combining methods take the form of a linear combination of the individual forecasts:

$$(2) \qquad \hat{y}^h_{c,t+h|t} = w_{0,t} + \sum_{i=1}^{n} w_{i,t} \hat{y}^h_{i,t+h|t},$$

where $\hat{y}^h_{c,t+h|t}$ is a given combination forecast whose weights, $\{w_{i,t}\}^n_{i=0}$, are typically calculated using the individual out-of-sample forecasts and y^h_{t+h} observations available from the start of the holdout out-of-sample period to time t. For each of the combining methods, we form combination forecasts over the post-holdout out-of-sample period, yielding $\{\hat{y}^h_{c,t+h|t}\}^{T-h}_{t=R+P_0}$, for a total of $T - (h - 1) - (R + P_0)$ forecasts available for evaluation. We compare the forecasts generated by each of the 20 combining methods, as well as the AR benchmark model, with the actual observations of employment growth over the post-holdout out-of-sample period, $\{y^h_{t+h}\}^{T-h}_{t=R+P_0}$.[2]

2.2.1 Simple Combining Methods. We consider three simple methods of combining individual forecasts: mean, median, and trimmed mean. Stock and Watson (2003, 2004) find that simple

[1] Although data availability is not an issue for financial variables, some nonfinancial variables are only available after a 1- to 2-month lag. Given this, it will generally be infeasible to use the procedure described in the text in real time at horizons of 1 to 2 months. At horizons beyond 2 months, say, 5 months, it is feasible to use the procedure to generate a forecast of cumulative employment growth over the previous 2 and subsequent 3 months.

[2] To be clear, out-of-sample forecasts are generated for the individual ARDL models over the entire out-of-sample period, which consists of both the holdout and post-holdout periods, using the recursive procedure described in Section 2.1.

combining methods work well in forecasting inflation and output growth using a large number of potential predictors in the G7 countries. The mean sets $w_{0,t} = 0$ and $w_{i,t} = (1/n)$ for all i in equation (2); the median uses the sample median of $\{\hat{y}^h_{i,t+h|t}\}^n_{i=1}$; the trimmed mean sets $w_{0,t} = 0$ and $w_{i,t} = 0$ for the individual models that produce the smallest and largest forecasts at time t, while $w_{i,t} = 1/(n-2)$ for the remaining individual models.[3]

2.2.2 OLS Combining Methods. Granger and Ramanathan (1984) recommend combining forecasts using unrestricted OLS. We consider OLS combination forecasts where the OLS coefficients are estimated using either a recursive or rolling window. To compute the initial OLS combination forecast (for $y^h_{R+P_0+h}$) using a recursive window, we regress $\{y^h_{s+h}\}^{R+(P_0-1)-(h-1)}_{s=R}$ on a constant and $\{\hat{y}^h_{i,s+h|s}\}^{R+(P_0-1)-(h-1)}_{s=R}$, $i = 1,\ldots,n$, and set the combining weights in equation (2) equal to the estimated OLS coefficients. To construct the second combination forecast (for $y^h_{R+(P_0+1)+h}$), the OLS coefficients are estimated by regressing $\{y^h_{s+h}\}^{R+(P_0-1)-(h-1)+1}_{s=R}$ on a constant and $\{\hat{y}^h_{i,s+h|s}\}^{R+(P_0-1)-(h-1)+1}_{s=R}$, $i = 1,\ldots,n$, and the fitted OLS coefficients again serve as the combining weights in equation (2). We proceed in this fashion through the end of the available out-of-sample period. The OLS combination forecasts based on a rolling window are computed in a similar manner, with the exception that in computing the second combination forecast, for example, the OLS coefficients that serve as the combining weights in equation (2) are estimated by regressing $\{y^h_{s+h}\}^{R+(P_0-1)-(h-1)+1}_{s=R+1}$ on a constant and $\{\hat{y}^h_{i,s+h|s}\}^{R+(P_0-1)-(h-1)+1}_{s=R+1}$, $i = 1,\ldots,n$.

2.2.3 WLS Combining Methods. Diebold and Pauly (1987) argue that combination forecasts based on time-varying weights can enhance forecasting performance in the presence of structural change. We use their "t-lambda" method. It follows the OLS combining method based on a recursive estimation window described in Section 2.2.2 above, with the exception that the combining weights are calculated using WLS instead of OLS. Diebold and Pauly (1987) recommend the weighting matrix $\Psi = diag[\psi_{tt}] = kt^{\lambda}$, where $k, \lambda > 0$, t = $1,\ldots,T$, and T is the number of observations used in the WLS regression. Under this approach, observations from the recent past receive more weight than observations from the distant past when computing the combining coefficients.[4] We consider $\lambda = 1$, which corresponds to weights that decrease at a constant rate as we move further into the past, and $\lambda = 3$, which corresponds to weights that decrease at an increasing rate.

2.2.4 Discount MSFE Combining Methods. Stock and Watson (2004) consider a combining method, where the weights in equation (2) depend inversely on the historical forecasting performance of the individual models. Their discount (or inverse) MSFE combining method employs the weights,

$$w_{i,t} = m_{it}^{-1} / \sum_{j=1}^{n} m_{j,t}^{-1}, \tag{3}$$

where

$$m_{i,t} = \sum_{s=R}^{t-h} \delta^{t-h-s}\left(y^h_{s+h} - \hat{y}^h_{i,s+h|s}\right)^2, \tag{4}$$

$w_{0,t} = 0$, and δ is a discount factor. Note that when $\delta = 1$, there is no discounting and equation (3) yields the optimal combination forecast derived by Bates and Granger (1969) for the case where the individual forecasts are uncorrelated; when $\delta < 1$, greater importance is attached to the recent forecasting performance of the individual models. We consider δ values of 1.0 and 0.9. Stock and Watson (2004) also consider a "most recently best" approach, where the "combination" forecast is the forecast corresponding to the individual model with the best forecasting performance over the previous year, and we include this approach in our analysis.

2.2.5 Bayesian Shrinkage Methods. In the presence of a relatively large number of individual forecasts, Bayesian shrinkage techniques may be helpful in forming combination forecasts, as suggested by Clemen and Winkler (1986) and

[3] The simple combining methods obviously do not require holdout out-of-sample observations.

[4] Using familiar notation, the WLS estimator can be expressed as $\hat{\beta}_{WLS} = (X'\Psi^{-1}X)^{-1}(X'\Psi^{-1}Y)$. Note that the value of k is arbitrary, because it disappears in the computation of the WLS estimator.

Diebold and Pauly (1990). We follow Stock and Watson (2004) and consider the following shrinkage combination forecast, which Diebold and Pauly (1990) show can be viewed as a Bayesian estimator:

$$w_{i,t} = \lambda\hat{\beta}_{i,t} + (1-\lambda)(1/n), \tag{5}$$

where $w_{0,t} = 0$, $\hat{\beta}_{i,t}$ is the OLS coefficient estimate corresponding to individual forecast i (the OLS coefficients are estimated using the recursive window scheme described in Section 2.2.2 above, with the exception that the intercept term is restricted to zero), $\lambda = \max\{0,1 - \kappa[n/(t - h - R - n)]\}$, and κ is a parameter that governs the degree of shrinkage toward equal weights. Larger values of κ correspond to smaller values of λ and thus more shrinkage toward equal weights. We consider κ values of 0.5 and 1.0.

2.2.6 Cluster Combining Methods. Aiolfi and Timmermann (2005) investigate persistence in forecasting performance and develop conditional combining methods. We use their $C(K, PB)$ algorithm, which proceeds as follows. To form the initial combination forecast, we first compute the MSFE for the individual forecasts $\{\hat{y}^h_{i,s+h|s}\}_{s=R}^{R+(P_0-1)-(h-1)}$, $i = 1,\ldots,n$, and group the individual models into K equal-sized clusters, where the first cluster contains the individual models with the lowest MSFE values, the second cluster contains the individual models with the next-lowest MSFE values, and so on. The first combination forecast, $\hat{y}^h_{c,R+P_0+h}$, is the average of the individual forecasts of $y^h_{R+P_0+h}$ generated by the models included in the first cluster. To form the second combination, we compute the MSFE for the individual forecasts $\{\hat{y}^h_{i,s+h|s}\}_{s=R+1}^{R+(P_0-1)-(h-1)+1}$, $i = 1,\ldots,n$, and group the individual models into clusters (so that the clusters are formed based on a rolling window), and the second combination forecast, $\hat{y}^h_{c,R+(P_0+1)+h}$, is the average of the individual forecasts of $y^h_{R+(P_0+1)+h}$ included in the first cluster. We proceed in this manner through the end of the available out-of-sample period. Following Aiolfi and Timmermann (2005), we consider $K = 2$ and $K = 3$ in our applications.

2.2.7 Model Selection Combining Methods. Swanson and Zeng (2001) consider combining methods based on model selection. We use their M-TST model selection approach, which uses a general-to-specific modeling procedure. We proceed as described in Section 2.2.2 above for the OLS combining method based on a rolling window, with the exception that we first examine the t-statistics corresponding to the estimated slope coefficients of the combining regression, where the t-statistics are calculated using heteroskedasticity and autocorrelation consistent (HAC) standard errors.[5] If any of the individual t-statistics are less then 1.645 in absolute value, we exclude these individual forecasts from the OLS regression used to estimate the combining weights. If all of the t-statistics are less than 1.645 in absolute value, we include all of the individual forecasts in the OLS regression.[6]

2.2.8 Principal Component Combining Methods. Chan, Stock, and Watson (1999) and Stock and Watson (2004) consider forming combination forecasts using the first m principal components of the individual forecasts. Let $\hat{F}^h_{1,s+h|s},\ldots,\hat{F}^h_{m,s+h|s}$, $s = R,\ldots,t$, represent the first m estimated principal components of the uncentered second-moment matrix of the individual forecasts, $\hat{y}^h_{i,s+h|s}$, $i = 1,\ldots,n$, $s = R,\ldots,t$. To form a combination forecast of y^h_{t+h} based on the fitted principal components, we estimate the regression model,

$$y^h_{s+h} = \phi_1\hat{F}^h_{1,s+h|s} + \ldots + \phi_m\hat{F}^h_{m,s+h|s} + v^h_{s+h}, \tag{6}$$

where $s = R,\ldots,t - h$. The combination forecast is given by $\hat{y}^h_{c,t+h|t} = \hat{\phi}_1\hat{F}^h_{1,t+h|t} + \ldots + \hat{\phi}_m\hat{F}^h_{m,t+h|t}$, where $\hat{\phi}_1,\ldots,\hat{\phi}_m$ are the OLS estimates of $\phi_1,\ldots,\phi_m$, respectively, in equation (6). We use $m = 1$ and $m = 2$ in computing forecasts using the principal component (PC) method.

2.2.9 Approximate Bayesian Model Averaging Combining Methods. Following Garratt et al. (2003), we compute combining weights using approximate Bayesian model averaging (ABMA),

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[5] We use Newey and West (1987) HAC standard errors with a lag truncation of $h - 1$.

[6] Swanson and Zeng (2001) also consider model selection based on the AIC or Schwarz information criterion (SIC). However, this involves computing the AIC or SIC for every possible combination of individual forecasts in the OLS regression model, which is impractical in our applications, as $n = 22$ so that there are $2^n - 1 = 4{,}194{,}303$ possible combinations of the individual forecasts.

in which functions of the SIC are used to approximate the posterior probabilities of the individual models (Draper, 1995). The combining weights can be expressed as

$$(7)\qquad w_{i,t} = e^{\Delta_{i,t}} \Big/ \sum_{j=1}^{n} e^{\Delta_{j,t}},\ \ i = 1,\ldots,n,$$

where $\Delta_{i,t} = SIC_{i,t}^{h} - \max_j(SIC_{j,t}^{h})$, and $SIC_{i,t}^{h}$ is the SIC corresponding to the fitted ARDL model i given by equation (1) used to generate $\hat{y}_{i,t+h|t}^{h}$. Garratt et al. (2003) also follow Burnham and Anderson (1998) and compute weights using the AIC, so that $\Delta_{i,t} = AIC_{i,t}^{h} - \max_j(AIC_{j,t}^{h})$ in equation (7). We consider ABMA combining weights based on both the SIC and AIC.[7]

2.2.10 Exponential Reweighting Combining Methods. Yang (2004) develops what he labels the AFTER (aggregated forecast through exponential reweighting) algorithm to combine forecasts from individual models. Yang (2004) shows that the algorithm can be viewed as an optimal combination procedure under fairly general conditions. The weights for the AFTER algorithm are given by

$$(8)\qquad w_{i,t} = \theta_{i,t} \Big/ \sum\nolimits_{j=1}^{n} \theta_{j,t},$$

where

$$(9)\qquad \theta_{i,t} = \prod_{s=R}^{t-h} \hat{\upsilon}_{i,s}^{0.5} e^{-0.5 \sum_{s=R}^{t-h} [(y_{s+h}^{h} - \hat{y}_{i,s+h|s}^{h})^2 / \hat{\upsilon}_{i,s}]},$$

and $\hat{\upsilon}_{i,t}$ is the OLS estimate of the variance of ε_{t+h}^{h} for the fitted ARDL forecasting model i (equation (1)) used to generate $\hat{y}_{i,t+h|t}^{h}$.

3 EMPIRICAL RESULTS

3.1 Data

Missouri employment growth is measured as the first difference in the log-levels of seasonally adjusted Missouri employment (multiplied by 100). The cumulative Missouri employment growth is divided by h, thereby expressing employment growth at the average monthly rate over the forecast horizon. We consider 9 regional and 13 national economic variables, for a total of 22 potential predictors ($x_{i,t}$ in equation (1)), which include labor market, production, and financial variables that are commonly used to forecast economic activity. The 9 regional variables are the Missouri unemployment rate and employment growth in the eight states that border Missouri (Arkansas, Illinois, Iowa, Kansas, Kentucky, Nebraska, Tennessee, and Oklahoma). The data are seasonally adjusted and, like the Missouri employment data, are from the Bureau of Labor Statistics (BLS). Based on availability, the data span from 1976:01 to 2005:01.[8] The 13 national economic variables are the following: U.S. employment; U.S. unemployment rate; capacity utilization rate; average weekly manufacturing hours; average weekly initial claims for unemployment insurance; manufacturers' new orders for consumer goods and materials; vendor performance (sales); manufacturers' new orders for nondefense capital goods; building permits; stock prices (S&P 500 index); interest rate spread (10-year Treasury bond yield minus the federal funds rate); national overtime; and industrial production. The U.S. employment and unemployment rate series are from the BLS; the remaining national variables are from the Conference Board. With three exceptions, all of the national variables are

[7] Like the simple combining methods, the ABMA combining methods do not require holdout out-of-sample observations.

[8] Supporting our specifications, the unit root tests of Ng and Perron (2001) clearly indicate that the log-levels of Missouri employment and employment in the eight bordering states are $I(1)$, while the Missouri unemployment rate is $I(0)$. We also tested for cointegration between the log-levels of Missouri employment and the log-levels of employment in each of the eight bordering states because equation (1) should potentially include an error-correction term in the event of cointegration. We only found evidence of cointegration between the log-levels of employment in Missouri and Kansas and Missouri and Oklahoma. However, Missouri employment appears weakly exogenous with respect to both of these variables, so the inclusion of an error-correction term in equation (1) is not necessary.

[9] With one exception, Ng and Perron (2001) unit root tests clearly support our specifications for the national variables. The exception is the U.S. unemployment rate, where the unit root null hypothesis cannot be rejected at the 10 percent level; however, the null hypothesis is very nearly rejected at the 10 percent level. We obtain similar results in our applications below if we the use first differences of the U.S. unemployment rate instead of the levels. Where relevant, we also found little evidence that an error-correction term needs to be included in equation (1) when any of the national variables serve as predictors.

Table 1

MSFE statistics for the Individual ARDL Models of Missouri Employment Growth, 1995:01–2005:01 Out-of-Sample Period

Variable	*h* = 3	*h* = 6	*h* = 12	*h* = 24
AR benchmark	1.92	1.21	0.88	0.84
Regional variables				
Missouri unemployment rate	0.97	0.94	1.01	0.98
Arkansas employment	1.00	1.00	0.92	0.93
Illinois employment	1.04	0.99	0.94	1.18
Iowa employment	0.98	0.92	0.95	0.91
Kansas employment	0.95	0.91	0.92	1.32
Kentucky employment	**0.90**	0.86	0.79	0.87
Nebraska employment	1.17	1.08	1.04	0.95
Tennessee employment	0.94	1.02	**0.70**	**0.76**
Oklahoma employment	0.97	1.00	1.06	1.22
National variables				
U.S. employment	1.11	0.93	0.98	1.16
U.S. unemployment rate	1.47	1.63	2.14	2.18
Capacity utilization	1.07	1.19	1.09	1.08
Average weekly hours, manufacturing	1.28	1.05	0.94	1.00
Unemployment claims	1.15	1.05	0.94	0.91
New manufacturing orders	1.06	1.03	0.96	0.98
Vendor sales	1.00	1.00	1.00	1.00
New manufacturing capital orders	1.08	1.15	1.26	1.30
Building permits	1.02	1.03	0.96	0.99
Stock market index	1.41	1.90	2.17	1.54
Interest rate spread	1.09	1.01	1.07	0.92
National overtime	0.95	**0.83**	0.87	0.94
Industrial production	1.30	0.98	1.00	1.00

NOTE: The first row reports the MSFE for the AR benchmark model; the remaining rows report the ratio of the MSFE for the individual ARDL model to the MSFE for the AR benchmark model. A bold entry signifies the ARDL model with the lowest MSFE at a given horizon.

measured in monthly growth rates (first differences of log-levels multiplied by 100); the three exceptions are the U.S. unemployment rate, capacity utilization rate, and interest rate spread, which are specified in levels.[9] While we do not claim that our list of 9 regional and 13 national variables constitutes an exhaustive list of potential predictors of Missouri employment growth, it does include a large number of potentially relevant predictors that are likely to be useful for our analysis.

3.2 Out-of-Sample Forecasting Results

We evaluate out-of-sample forecasts of Missouri employment growth over the 1995:01 to 2005:01 period. This period includes the late-1990s expansion, 2001 recession, and subsequent "jobless" recovery—an informative period in which to evaluate forecasts of Missouri employment growth. We consider "short" forecast horizons of 3 and 6 months and "long" forecast horizons of 12 and 24 months. As discussed in Section 2.2

Table 2

Forecast Combining Results for Missouri Employment Growth, 1995:01–2005:01 Out-of-Sample Period

	h = 3				*h* = 6			
Combination method	MSFE	$\hat{\alpha}_0$	$\hat{\alpha}_1$	R^2	MSFE	$\hat{\alpha}_0$	$\hat{\alpha}_1$	R^2
AR benchmark	1.92	0.08	0.57*	0.08	1.21	–0.08	0.65	0.10
Mean	0.96	–0.02	0.67	0.10	0.92	–0.60	1.14	0.21
Median	0.94	–0.03	0.68	0.10	0.92	–0.54	1.07	0.22
Trimmed mean	**0.93**	–0.06	0.72	0.11	**0.90**	–0.57	1.12	0.23
OLS, recursive	1.44	0.53*	0.17**	0.02	1.56	0.33	0.35**	0.20
OLS, rolling	2.28	0.61**	0.07**	0.01	2.45	0.31	0.25**	0.21
WLS: t-lambda, $\lambda = 1$	1.65	0.58*	0.11**	0.01	1.90	0.36	0.28**	0.20
WLS: t-lambda, $\lambda = 3$	2.06	0.58*	0.10**	0.01	2.30	0.35	0.24**	0.20
Discount MSFE, $\delta = 1.0$	0.95	–0.02	0.67	0.10	0.92	–0.64	1.16	0.22
Discount MSFE, $\delta = 0.9$	0.95	0.01	0.65	0.10	0.91	–0.53	1.06	0.18
Most recently best	1.07	0.20	0.46**	0.10	1.03	0.55	0.12**	0.01
Shrinkage, $\kappa = 0.5$	1.18	0.47	0.25**	0.03	1.21	0.28	0.48**	0.24
Shrinkage, $\kappa = 1.0$	1.08	0.38	0.35**	0.04	1.02	0.19	0.59	0.24
Cluster: $C(2, PB)$	0.96	–0.02	0.67	0.09	0.94	–0.54	1.01	0.16
Cluster: $C(3, PB)$	0.96	0.07	0.61	0.08	0.94	–0.70	1.14	0.19
Model selection: M-TST	1.82	0.57*	0.12**	0.02	1.93	0.32	0.22**	0.13
PC, $m = 1$	0.96	0.00	0.65	0.10	0.97	–0.68	1.04	**0.30**
PC, $m = 2$	0.96	–0.07	0.71	0.09	0.97	–0.62	0.96	0.25
ABMA, SIC	1.07	0.21	0.45**	0.06	1.01	0.05	0.55	0.05
ABMA, AIC	1.05	0.13	0.50**	0.09	1.17	–0.26	0.77	0.23
AFTER	0.94	–0.05	0.69	**0.12**	0.99	–0.14	0.71	0.19

NOTE: The first row reports the MSFE for the AR benchmark model; the remaining rows report the ratio of the MSFE for the combining method to the MSFE for the AR benchmark model. $\hat{\alpha}_0$, $\hat{\alpha}_1$, and R^2 are the intercept estimate, slope estimate, and goodness-of-fit measure, respectively, for the MZ regression. A bold entry signifies the combining method with the lowest MSFE or the highest R^2 at a given horizon; * and ** indicate significance at the 5 percent and 1 percent levels, respectively, for a test of the null hypothesis that $\alpha_0 = 0$ ($\alpha_1 = 1$) for $\hat{\alpha}_0$ ($\hat{\alpha}_1$).

above, we need a holdout out-of-sample period in order to compute most of the combination forecasts, and we use the 60 observations preceding 1995:01 as the holdout out-of-sample period.

Table 1 reports out-of-sample forecasting results for the AR benchmark model and the individual ARDL models. The table reports the MSFE statistics for the AR benchmark model and the ratio of the MSFE for the individual ARDL models to the MSFE for the AR benchmark model. The first row of the table shows that the MSFE declines as the horizon increases for the AR benchmark model, suggesting that more accurate forecasts of average monthly Missouri employment growth are available at longer horizons. Observe that either five or six of the nine individual ARDL models based on the regional variables have lower MSFE statistics than those for the AR benchmark model at all reported horizons. The performance of national variables is poorer, as most individual ARDL models do not have lower MSFE statistics than those for the AR benchmark, particularly at shorter horizons. It would seem difficult to ascertain a priori which of the individual potential

h = 12				h = 24			
MSFE	$\hat{\alpha}_0$	$\hat{\alpha}_1$	R^2	MSFE	$\hat{\alpha}_0$	$\hat{\alpha}_1$	R^2
0.88	–0.08	0.65	0.10	0.84	–0.18	0.63	0.04
0.84	–0.60	1.14	0.21	0.83	–0.39	0.88	0.05
0.85	–0.54	1.07	0.22	0.94	–0.49	0.90	0.07
0.82	–0.57	1.12	0.23	0.83	–0.36	0.85	0.07
1.27	0.33	0.35**	0.20	0.95	–0.12	0.58	0.52
2.81	0.31	0.25**	0.21	1.62	0.14	0.36	0.47
1.87	0.36	0.28**	0.20	1.30	–0.08	0.50	0.50
2.65	0.35	0.24**	0.20	1.94	–0.05	0.42**	**0.53**
0.85	–0.64	1.16	0.22	0.91	–0.72	1.10	0.07
0.87	–0.53	1.06	0.18	0.91	–0.68	1.08	0.05
1.32	0.55	0.12**	0.01	0.87	–0.09	0.60	0.15
0.85	0.28	0.48**	0.24	**0.51**	–0.16	0.80	0.33
0.71	0.19	0.59	0.24	0.53	–0.28	0.92	0.28
0.97	–0.54	1.01	0.16	1.06	0.54	0.07	0.00
0.96	–0.70	1.14	0.19	1.11	0.34	0.22	0.00
3.00	0.32	0.22**	0.13	1.73	0.13	0.36**	0.48
1.00	–0.68	1.04	**0.30**	1.36	–1.15	1.17	0.23
1.10	–0.62	0.96	0.25	1.27	–0.43	0.73	0.12
1.00	0.05	0.55	0.05	1.16	0.99**	–0.32**	0.04
0.93	–0.26	0.77	0.23	0.90	–0.95	1.28	0.13
0.92	–0.14	0.71	0.19	0.89	–0.89	1.23	0.15

predictors in the ARDL models are likely to display the best forecasting ability, and this provides a motivation for considering methods for combining the large number of individual ARDL forecasts.

Table 2 presents results for the 20 forecast combining methods over the 1995:01–2005:01 out-of-sample period. Similar to Table 1, Table 2 reports the MSFE for the AR benchmark model and the ratio of the MSFE for a given combining method to the MSFE for the AR benchmark model. Table 2 also reports the estimated intercept, estimated slope, and R^2 statistic for a Mincer and Zarnowitz (MZ, 1969) regression of the form

$$y_{t+h}^{h} = a_0 + a_1 \hat{y}_{c,t+h|h}^{h} + \eta_{t+h}^{h}, \tag{10}$$

where $a_0 = 0$ and $a_1 = 1$ when the forecasts are unbiased. We indicate in Table 2 whether $\hat{a}_0$ ($\hat{a}_1$) is significantly different from 0 (1), where $\hat{a}_0$ ($\hat{a}_1$) is the OLS estimate of a_0 (a_1) in equation (10).[10]

[10] The *t*-statistics used to assess the statistical significance are based on Newey and West (1987) standard errors with a lag truncation of h–1.

Given that the forecasts are unbiased, the R^2 statistic provides a measure of the ability of the forecasts to explain movements in actual Missouri employment growth.

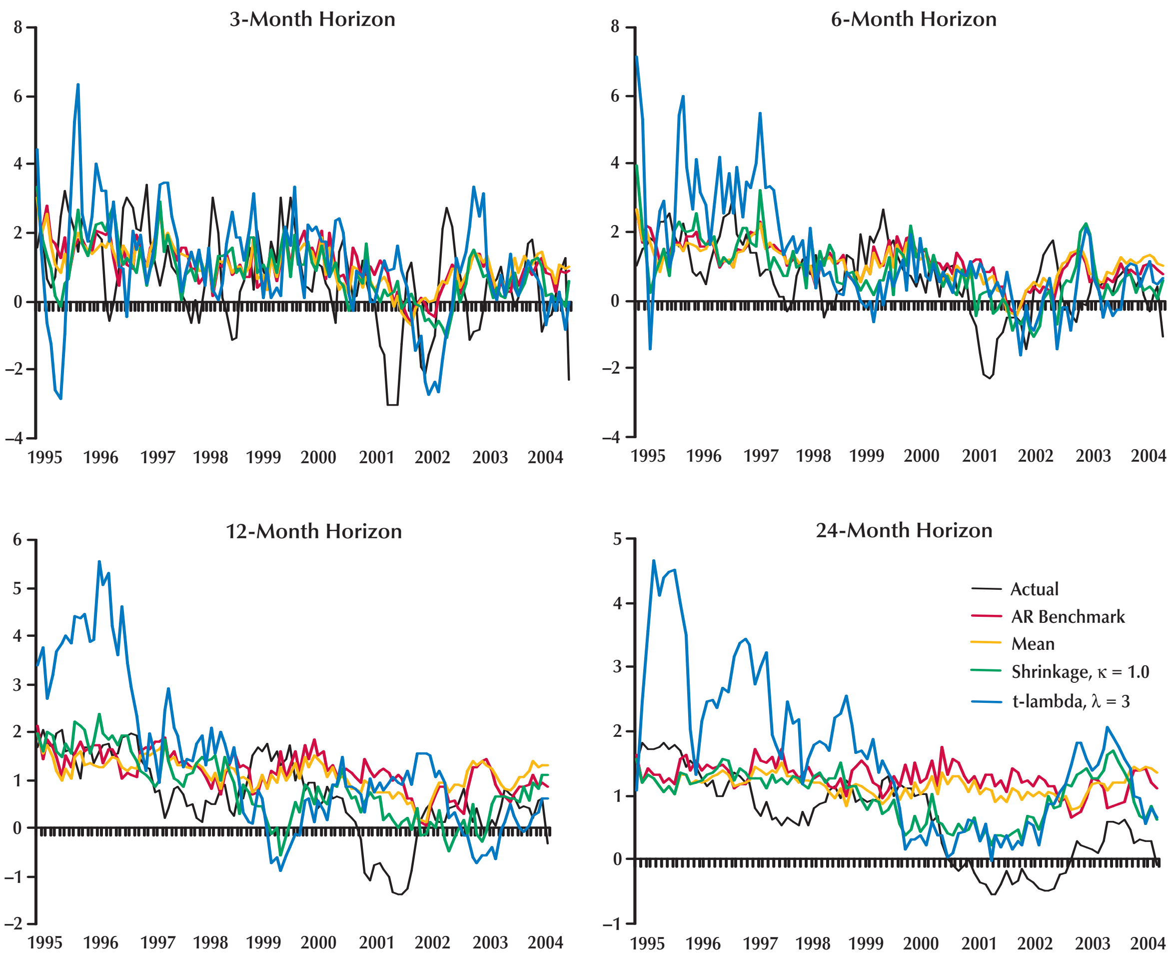
Among the forecast combining methods, there is considerable dispersion of results. The simple combining methods (mean, median, and trimmed mean) consistently outperform the AR benchmark, with reductions in MSFE of around 5 percent to 15 percent relative to the AR benchmark, and the R^2 statistic for the MZ regressions are greater than those for the AR benchmark. In addition, the estimated intercept and slope coefficients are not significantly different from 0 and 1, respectively, so that the simple combining methods appear to produce unbiased forecasts. At shorter horizons of 3 and 6 months, the trimmed mean combining method outperforms all of the other combining methods in terms of MSFE. In terms of the R^2 statistic of the MZ regression, the AFTER procedure performs marginally better than other simple combining methods at $h = 3$, and the PC ($m = 1$) method provides the best MZ fit at $h = 6$, with an R^2 statistic of 0.30. Both the AFTER and PC methods have estimated intercept and slope coefficients that are not significantly different from 0 and 1, respectively, in the MZ regression. Observe that many of the combining methods, especially the OLS and t-lambda methods, produce forecasts that are both substantially less accurate than the AR benchmark and biased according to the MZ regression results. Also note that the trimmed mean, the best performing combining method at horizons of 3 and 6 months, has an MSFE that is reasonably close to that of the best performing individual ARDL model at these horizons (Kentucky employment and national overtime, respectively). Given that it will be very difficult a priori for a researcher to select the individual variable that will perform the best, this helps to demonstrate the usefulness of the trimmed mean combining method at shorter horizons.

At longer horizons of 12 and 24 months, Table 2 shows that several of the combining methods lead to sizable reductions in MSFE and increases in the MZ R^2 statistics relative to the AR benchmark model. In particular, the shrinkage method with $\kappa = 1.0$ ($\kappa = 0.5$) leads to reductions in MSFE of 29 percent (15 percent) and 47 percent (49 percent) relative to the AR benchmark at horizons of 12 and 24 months, respectively. The coefficient estimates from the MZ regression indicate that the shrinkage forecasts are unbiased, with the exception of the slope coefficient when $\kappa = 0.5$ and $h = 12$. In addition, the shrinkage method with $\kappa = 1.0$ performs nearly as well as the best individual ARDL model (Tennessee employment) at the 12-month horizon and better than the best individual ARDL model (Tennessee employment) at the 24-month horizon, further demonstrating the usefulness of the shrinkage method with $\kappa = 1.0$ at longer horizons. The OLS and t-lambda procedures yield relatively high MZ R^2 statistics at $h = 12$ and $h = 24$, but the estimated slope coefficients in the MZ regressions are significantly less than 1 at the 1-year horizon. Moreover, these combining methods have MSFE statistics well above those for the AR benchmark, with the exception of the OLS recursive method when $h = 24$. The discount MSFE method appears to offer reasonably large reduction in MSFE relative to the AR benchmark model at horizons of 12 and 24 months. However, the reductions in MSFE associated with the discount MSFE methods are smaller than those for the simple combining methods. In fact, the simple combining methods offer fairly sizable reductions in MSFE relative to the AR benchmark and generate unbiased forecasts at horizons of 12 and 24 months. Overall, the simple combining methods perform consistently well at all reported horizons in Table 2 and seem to offer a low-cost way of generating reliable forecasts of Missouri employment growth.

To gain further insight into the relative forecasting performances of some of the combining methods, Figure 1 plots the realized observations of y_{t+h}^{h} and the forecasts generated by the AR benchmark model and the mean; shrinkage, $\kappa = 1.0$; and t-lambda, $\lambda = 3$ combining methods. The shrinkage, $\kappa = 1.0$ method is shown because it has the lowest MSFE at $h = 12$ and next-to-lowest MSFE at $h = 24$; the t-lambda, $\lambda = 3$ method is shown because it has the highest MZ R^2 statistic at $h = 24$. A problem with the t-lambda method,

Figure 1

Actual Missouri Employment Growth and Select Forecasts, 1995:01–2005:01 Out-of-Sample Period



especially at horizons of 6, 12, and 24 months, is that it is much more volatile than the actual realizations. Despite the fact that the t-lambda method has relatively high MZ R^2 statistics at horizons of 6, 12, and 24 months, the overly volatile nature of the t-lambda forecasts causes its MSFE to be substantially greater than that of the AR benchmark, and the estimated slope coefficient in the MZ regression is consistently significantly less than 1. At horizons of 12 and 24 months, Figure 1 shows that the shrinkage method tends to do the best job of tracking the decline in Missouri employment growth associated with the 2001 recession. Given that turning points are notoriously difficult to predict, this suggests that the shrinkage method forecasts are quite useful at longer horizons. We also see from Figure 1 that the mean generally does a better job than the AR benchmark model at tracking Missouri employment growth at all horizons. The mean forecasts are less volatile than the shrinkage and t-lambda forecasts, so they provide more reliable forecasts at shorter horizons.

Table 3

Shrinkage, κ = 1.0 Combining Weights for the Individual ARDL Model Forecasts of Missouri Employment Growth for Select Months, 1995:01–2005:01 Out-of-Sample Period, h = 12

Variable	1996:01	1997:01	1998:01	1999:01	2000:01	2001:01	2002:01	2003:01	2004:01
Regional variables									
Missouri unemployment rate	0.01	0.20	0.21	0.38	0.23	–0.18	–0.03	–0.02	–0.01
Arkansas employment	0.17	0.19	0.18	0.16	0.17	0.21	0.30	0.28	0.26
Illinois employment	0.02	–0.09	–0.10	–0.14	–0.05	0.02	0.06	0.01	–0.04
Iowa employment	0.28	0.25	0.28	0.25	0.26	0.31	0.28	0.26	0.25
Kansas employment	0.39	0.31	0.04	–0.16	–0.34	–0.17	–0.19	–0.06	0.00
Kentucky employment	0.00	0.27	0.23	0.11	0.23	0.37	**0.59**	**0.88**	**0.84**
Nebraska employment	**0.51**	**0.58**	**0.66**	**0.67**	**0.60**	0.45	0.43	0.40	0.35
Tennessee employment	0.01	–0.02	–0.08	–0.03	0.18	0.17	0.21	0.34	0.42
Oklahoma employment	0.17	0.20	0.21	0.38	0.23	–0.18	–0.03	–0.02	–0.01
National variables									
U.S. employment	–0.55	–0.68	–0.64	–0.68	–0.66	–0.58	–1.05	–1.27	–1.25
U.S. unemployment rate	–0.12	–0.29	–0.22	–0.23	–0.12	–0.12	–0.20	–0.26	–0.19
Capacity utilization	–0.22	–0.35	–0.45	–0.68	–0.60	–0.30	–0.47	–0.44	–0.35
Average weekly hours, manufacturing	–0.22	–0.14	0.06	0.08	0.12	0.02	0.02	0.00	–0.05
Unemployment claims	0.23	0.16	0.23	0.19	0.25	0.26	0.45	0.43	0.35
New manufacturing orders	0.02	–0.08	–0.09	–0.12	–0.11	–0.19	–0.22	–0.23	–0.26
Vendor sales	0.36	0.56	0.58	0.55	0.57	**0.58**	0.52	0.53	0.46
New manufacturing capital orders	0.03	0.06	0.02	0.05	–0.03	–0.07	–0.14	–0.22	–0.23
Building permits	0.13	0.07	0.06	0.10	0.05	0.01	0.01	–0.07	–0.12
Stock market index	–0.51	–0.78	–0.78	–0.62	–0.51	–0.42	–0.32	–0.28	–0.20
Interest rate spread	0.06	0.18	0.17	0.32	0.34	0.24	0.40	0.46	0.48
National overtime	0.12	0.21	0.12	0.09	–0.14	–0.15	–0.15	–0.17	–0.13
Industrial production	0.27	0.30	0.34	0.33	0.31	0.41	0.27	0.22	0.20

NOTE: A bold entry signifies the ARDL model that receives the highest weight at the given date; 0.00 indicates <0.005.

However, the mean forecasts are too smooth at longer horizons and thus are not as adept as the shrinkage forecasts in tracing swings in Missouri employment growth.

Given the relatively good forecasting performance of the shrinkage, $\kappa = 1.0$ method at longer horizons, it is interesting to examine the weights used to compute the combination forecasts for this method. Table 3 reports the weights associated with the individual forecasts using the shrinkage, $\kappa = 1.0$ method for the first month of most of the years of the forecast evaluation period. Each weight would be 0.045 under equal weighting. However, the weights on the individual forecasts often differ markedly from equal weighting in Table 3, and we also witness sizable changes in some of the weights from year to year. Several of the regional variables, such as Nebraska, Iowa, and Kentucky employment, are heavily positively weighted throughout most of the period, indicat-

Table 4

Forecast Combining Results for Missouri and Illinois Employment Growth, Alternative Out-of-Sample Periods, $h = 12$

Combination method	Missouri 1995:01–1999:12	Missouri 2000:01–2005:01	Illinois 1995:01–2005:01	Illinois 2000:01–2005:01
AR benchmark	0.47	1.41	1.72	3.22
Mean	0.88	0.83	0.81	0.81
Median	0.94	0.82	0.82	0.83
Trimmed mean	0.99	0.79	0.84	0.84
OLS, recursive	2.25	0.61	1.34	1.00
OLS, rolling	7.28	1.04	1.28	1.01
WLS: t-lambda, $\lambda = 1$	4.25	0.75	1.39	0.99
WLS: t-lambda, $\lambda = 3$	7.10	0.90	1.40	1.08
Discount MSFE, $\delta = 1.0$	0.89	0.84	0.82	0.80
Discount MSFE, $\delta = 0.9$	0.93	0.84	0.81	0.82
Most recently best	1.67	1.00	**0.59**	**0.53**
Shrinkage, $\kappa = 0.5$	1.25	**0.42**	0.90	0.64
Shrinkage, $\kappa = 1.0$	**0.88**	0.44	0.82	0.64
Cluster: $C(2, PB)$	0.97	0.90	0.89	0.78
Cluster: $C(3, PB)$	0.95	0.90	0.88	0.74
Model selection: M-TST	6.85	1.29	7.36	1.02
PC, $m = 1$	1.31	0.91	0.83	0.83
PC, $m = 2$	1.26	1.07	0.98	0.83
ABMA, SIC	1.04	1.00	1.41	1.39
ABMA, AIC	0.96	0.94	0.72	0.69
AFTER	0.98	0.84	1.06	1.03

NOTE: The first row reports the MSFE for the AR benchmark model; the remaining rows report the ratio of the MSFE for the combining method to the MSFE for the AR benchmark model. A bold entry signifies the combining method with the lowest MSFE.

ing the important role for some of the regional variables in forecasting Missouri employment growth.[11] Tennessee and Oklahoma employment are interesting in that the weights display significant increases and decreases, respectively, over the evaluation period. In terms of the national variables, vendor sales, industrial production, and the interest rate spread possess large positive weights throughout most of the period; other national variables such as U.S. employment, capacity utilization, and the stock market index have large negative weights throughout most of the period. Overall, the shrinkage method at the 1-year horizon appears able to identify the individual forecasts that are the most accurate and to accommodate changes in relative forecasting accuracy in computing the combining weights. It would be difficult a priori to identify the particular individual model or models with the best forecasting ability at different points in time, and the shrinkage method provides an a priori procedure to cull potentially useful information from a large number of individual models.

[11] For instance, if we were to exclude the eight neighboring state employment variables, the ratio of the MSFE for the shrinkage, $\kappa = 1.0$ procedure to the MSFE for the AR benchmark increases to 1.01 at the 12-month horizon, well above the 0.71 figure in Table 2.

3.3 Robustness Checks

Table 4 reports results for alternative forecast evaluation periods for Missouri employment growth and employment growth for a neighboring state, Illinois, at the 1-year horizon.[12] The results for Missouri show that the MSFE for the AR forecast procedure is more than three times lower for the 1995:01–1999:12 period than the 2000:01–2005:01 period. This reflects fairly sharp employment changes and potential data problems (for example, occasional contradictory reports between BLS monthly and quarterly reports on both the national and state levels) during the latter period.[13] Several of the combining methods in the earlier period produce very poor results, with MSFE ratios substantially above 1, while the shrinkage, $\kappa = 1.0$ method and mean combination forecast produce the lowest MSFE statistics. Both shrinkage forecasts produce the lowest MSFE during the latter period for Missouri. Overall, across our three evaluation periods (1995:01–2005:01, 1995:01–1999:01, 2000:01–2005:1), the shrinkage, $\kappa = 1.0$ method yields the lowest MSFE at the 1-year horizon for Missouri employment growth.

With respect to employment growth in Illinois, we consider the 1995:01–2005:01 and 2000:01–2005:01 forecast evaluation periods. The individual ARDL forecasts are based on the same set of national variables used for Missouri, as well as the Illinois unemployment rate and employment in the six states that border Illinois (Iowa, Indiana, Kentucky, Missouri, Michigan, and Wisconsin). The same time-series specifications given in Section 3.1 above are used for these variables. The most recently best method achieves the lowest MSFE for both evaluation periods, producing declines of 51 percent and 47 percent in MSFE relative to the AR benchmark. The simple combination forecasts outperform the AR benchmark over both evaluation periods, with declines in MSFE relative to the AR benchmark of 16 to 19 percent, and the shrinkage and ABMA, AIC methods also lead to sizable declines in MSFE relative to the AR benchmark, ranging from 10 to 36 percent. The results in the last two columns of Table 4 indicate that simple combining and shrinkage methods perform well with respect to forecasting employment growth at the 1-year horizon in both Missouri and Illinois.

4 CONCLUSION

There are a large number of potentially relevant predictors of regional economic variables such as Missouri employment growth. In this paper, we analyze 20 different methods from the extant literature for combining individual forecasts of Missouri employment growth from 1995:01 to 2005:01. The individual forecasts are generated by a large number of ARDL models based on potential predictors. Similar to Stock and Watson (2004), we find that simple combination methods work well, particularly at shorter forecast horizons of 3 and 6 months, and often outperform more complicated weighting procedures. At longer horizons of 12 and 24 months, we find that Bayesian shrinkage methods produce the most accurate forecasts, providing quite sizable reduction in MSFE relative to an AR benchmark model. The shrinkage combining methods also perform well at longer horizons over alternative evaluation periods and were able to track Missouri employment growth over the recent recession reasonably well. Examination of the combining weights used in the shrinkage methods indicates that a number of regional variables and a few national variables can play a significant role in improving forecasts of Missouri employment growth. Forecast combination procedures also lead to relatively large reductions in MSFE relative to an AR benchmark model when forecasting Illinois employment growth. Why do shrinkage combining methods work well in the present paper? Diebold and Lopez (1996, p. 256) offer helpful insight by observing that "the combining weights [of shrinkage combining methods] are coaxed toward the arithmetic mean, but the data are still allowed to speak, when (and if) the data have something to say." Shrinkage combining methods can thus take

[12] A 1-year horizon is typically important for state budgetary planning purposes.

[13] See Kliesen and Wall (2004) on reconciling the BLS jobless employment figures and Wall and Wheeler (2005) on large discrepancies in recent St. Louis employment.

advantage of the reliable performance of simple averaging across a diversity of variables while still allowing for particular variables to exert a stronger influence in certain situations. In future research, we plan to extend our analysis by developing combination forecasts of employment growth for each of the 50 individual U.S. states.

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