BUILDING A REGIONAL FORECASTING MODEL UTILIZING LONG-TERM RELATIONSHIPS AND SHORT-TERM INDICATORS

Keith R. Phillips

and

Chih-Ping Chang

June 1995

RESEARCH DEPARTMENT

WORKING PAPER

95-04

Federal Reserve Bank of Dallas

Building a Regional Forecasting Model Utilizing Long-term Relationships and Short-term Indicators[1]

Keith R. Phillips

Federal Reserve Bank of Dallas

Chih-Ping Chang

Southern Methodist University

June 1995

[1]The authors thank seminar participants of the November 1994 Federal Reserve Regional System Committee Meeting and Nathan S. Balke for helpful comments. The views expressed in this paper are those of the authors and should not be attributed to the Federal Reserve Bank of Dallas or the Federal Reserve System.

Abstract

Chang and Phillips develop a simple labor-demand error correction model of regional employment growth. The model is constructed to forecast well at both long-term and short-term horizons. In developing the model, we utilize past research which has found that relative nominal wages play an important role in explaining why some regions consistently grow faster than others. The variables in the model include regional employment, U.S. employment, an industry-adjusted relative wage measure, and a regional leading index. While the wage variable is used to capture long-term shifts in relative labor demand, the leading index is included to control for shorter-term cyclical shocks. Out-of-sample forecast errors from the model are shown to be smaller than errors from a model suggested by LeSage (1990a) which divides regional employment into base and nonbase and estimates a bivariate error-correction model.

I. Introduction

Regional economists are often asked to produce forecasts that range anywhere from several months to several years. Often, however, the forecasting technique and variables used by the analyst differ depending on whether the forecast is long- or short-term. For example, the analyst might use average weekly hours of manufacturing workers to forecast short-term cyclical movements in regional employment but would not likely use the same indicator for forecasting longer-term changes. In this article we attempt to build a simple regional forecasting model that utilizes short-term indicators and long-term relationships in an attempt to forecast well at both short-term and long-term forecasting horizons.

In a recent series of articles, LeSage (1990a, b) and Shoesmith (1995) developed several error-correction models to forecast regional employment. The purpose of the error-correction model is to determine and estimate any long-term equilibrium relationships that exist among nonstationary variables. Given that these stationary relationships can be established, using the Error-Correction Vector-Autoregressive Model (ECM) has more appeal than simply running models in first-differenced form and thus focusing on the short-term relationship between the variables. In this article we develop an ECM which utilizes the long-term equilibrium relationship between a region's relative employment and its relative wage. (The term *relative* refers to relative to the national average, unless stated otherwise.) The model also utilizes movements in a regional leading index in an attempt to capture upcoming short-term cyclical shocks.

We develop a four-variable ECM for the Texas economy and compute out-of-sample forecasts between March 1991 and March 1995 and for the 1986 Texas recession. Errors from one- to thirty-six-step ahead forecasts are compared to a two-variable ECM suggested by LeSage (1990a) which divides employment into base and nonbase. Our model performs better on average during each forecasting period and at almost all forecast horizons within each period. The results suggest that our model may be useful for forecasting total nonagricultural employment at the regional level.

II. Developing a Regional Forecasting Model

In an attempt to forecast employment for fifty different industries in Ohio, LeSage (1990b) finds significant cointegration between manhours, hourly earnings, and the consumer price index in seven industries. He then compares out-of-sample errors from an ECM for each of the fifty industries to errors from a differenced Vector-Autoregressive Model (VAR), two types of Bayesian VARs and a Bayesian ECM. In the seven industries with cointegration, he finds that the ECM performs best. He also finds, however, that the ECM performed well at the seven- to twelve-month forecast horizons, even in models that failed the cointegration tests.

In another paper, LeSage (1990a) forecasts total nonfarm employment for eight different metropolitan areas in Ohio using the ECM within the context of the economic base model. Using the three categories of durable, nondurable, and nonmanufacturing industries, LeSage defines base employment during any month as any positive residual of employment in an industry minus what the U.S. industry share would suggest. After dividing employment into base and nonbase, he finds evidence of cointegration between the two employment types in each of the eight cities. He then compares the results of the ECM to the four types of models described in LeSage (1990b) and to a forecast produced by an independent analyst. LeSage finds that the ECM and BECM used in the context of a dynamic economic base model performed the best out of the six models examined and concludes that the ECM used on base and nonbase employment was a simple, but effective, way to forecast regional employment.

The results of the two LeSage articles give strong evidence of the usefulness of the ECM in forecasting regional employment. The use of the economic base model, however, is less convincing. Although the economic base model has a long history in regional economics, in recent years it has faced increasing criticism.[2] While the LeSage results give evidence that base and nonbase industries share a long-term relationship and that growth in base industries leads to growth in nonbase industries, it may be possible to construct a simple regional ECM that has stronger links to regional employment growth.

The use of the ECM in regional forecasting is based on finding variables that share a long-term cointegrating relationship. Finding variables that share this type of relationship may often be difficult. For example, Hefner (1990) showed that for all eight Bureau of Economic Analysis regions, Gross State Product was not cointegrated with U.S. Gross Domestic Product, and Shoesmith (1992) showed that for every state except Vermont, U.S. nonfarm employment was not cointegrated with the state nonfarm employment. Shoesmith (1995) also showed that in his five-variable regional VAR model, only one out of four states tested had at least one significant cointegrating vector. These results emphasize the need for a strong theoretical notion of what variables are likely to experience stable long-term relationships.

One stylized fact important in long-run regional employment forecasting is that many regions consistently grow faster than others. Browne, Mieszkowski, and Syron (1980) in a study of the relative economic strength of the South, found that low nominal wages played an important role in attracting net investment into the South. Using surveys, other researchers-such as Nakosteen and Zimmer (1987) and Wheat (1986)-have confirmed the importance of relative nominal wages in regional job growth.

Kottman (1992) suggested that, for industries which export nationally, the decision on where to produce is a function of relative nominal wages, but for workers the decision on where

---

[2]For a good summary of the recent criticisms of the economic base model, see Krikelas (1992).

to work is a function of relative cost-of-living adjusted wages. Kottman points out that in terms of labor supply, "There is an emerging consensus in the literature that regional wage and income differentials have all but vanished once adjusted for cost-of-living and labor force characteristics." He concludes that the persistent migration of capital and labor appears to be almost solely due to labor demand - firms migrate to areas with low relative nominal wages. While workers' real wages are similar across regions, low nominal wages stimulate relatively strong job growth.

In our model, we adopt the view that the main factor driving relative regional employment growth is shifting labor demand which is driven mainly by relative nominal wages. As information flows into the high-wage regions and firms realize that they can increase profitability by moving to low-wage regions, firms begin to migrate. As firms migrate to the low-wage regions they build up the regional infrastructure, which motivates more firms to move in - gradually shifting out the labor demand curve in the low-wage regions. In the traditional labor-market model, shifts in labor demand result in a positive relationship between employment and real wages as the labor supply curve is upward sloping. In this model, low-wage regions experience relatively strong labor demand shifts, which result in stronger gains in wages and employment than in high-wage regions.[3] Thus, we expect a positive long-run relationship between a region's relative wage and its relative employment.

An industry is attracted to a region if wages in that region are relatively low. It is thus important to compare wages across the same industry and not just compare total manufacturing wages. We thus create a wage variable which measures the percentage difference in an industry's wage from the same industry in the nation at the two-digit Standard Industrial

[3]One implication of this model is that wages in the low-wage region would be driven up while wages in the high-wage region would be driven down - leading to wage convergence. Supporting this notion, Carlino and Mills (1993) find statistical evidence that per capita incomes have converged in the United States since 1929, and Browne (1989) points out that most of the movement in per capita personal income-at least over the past two decades-has been due to variations in wages.

Classification level. The wage differentials are weighted together by the industry's share of national manufacturing employment to form a composite measure of the region's relative manufacturing wage. Chart 1 shows that in Texas this wage measure appears to share a long-term positive relationship with Texas nonfarm employment.

While we expect a long-term relationship between relative wages and relative employment, we also expect employment to exhibit business-cycle patterns because of economic shocks such as oil price changes. To incorporate these short-term shocks, we include the Texas Leading Index which was designed to predict cyclical turning points in the state's economy.[4] The components of the Texas Leading Index are average weekly hours of production workers in manufacturing, an index of help-wanted advertising, an index of stock prices of companies based in Texas, new unemployment claims (inverted), real retail sales, permits to drill oil and gas wells, the real price of crude oil, the BEA national leading index, and an index of the real exchange rate of the countries Texas exports to (inverted). Phillips (1988, 1990) has shown the index to be a reliable indicator of turning points in the Texas economy with a lead time of three to eight months.

The long-run relationship suggested by the labor demand model is

$$\ln\left(\frac{TXEMP_t}{USEMP_t}\right)=\beta_0+\beta_1\ln(RW_t)+e_t \qquad \text{(RW)}$$

where ln is the natural log function, TXEMP and USEMP are the levels of nonfarm employment in Texas and the United States, and RW is a measure of Texas hourly wages relative to U.S. hourly wages. Preliminary testing showed that a significant long-term

[4]The use of regional leading indexes has proliferated over the past ten years. For a description of the typical index calculation and a listing of indexes available at the regional level, see Phillips (1994). For more information on the components in and the calculation of the Texas Leading Index, see Phillips (1990).

relationship holds between the two variables in this equation.[5] While this relationship is important in forecasting the relative growth of Texas employment in the long-run, it is also likely that changes in the Texas Leading Index (TXLI) are important in directly predicting shorter-term cyclical shocks to the Texas economy. However, since the TXLI is not likely to have proportional effects on both TXEMP and USEMP, we build the model based on the relationship between all four variables.

The four-equation ECM we estimate is

$$\Delta X_t = \mu + A(L)\Delta X_{t-1} + \alpha\beta X_{t-k-1} + \epsilon_t$$

where X_t is a vector containing four series - Texas' relative wage (rw), the Texas Leading Index (txli), Texas nonfarm employment (txemp), and U.S. nonfarm employment (usemp) - and Δ is the first-difference operator. The lowercase of a variable indicates its logarithmic value. A(L) is a lag polynomial of order k. βX is the error-correction term capturing the stationary long-run equilibrium among the four variables. For example, the equation of Texas nonfarm employment in the VAR model is specified as

$$\Delta txemp_t = \mu + \sum_{i=1}^{T} \lambda_i \Delta txemp_{t-i} + \sum_{i=1}^{T} \theta_i \Delta rw_{t-i} + \sum_{i=1}^{T} \gamma_i \Delta txli_{t-i} + \sum_{i=1}^{T} \eta_i \Delta usemp_{t-i}$$
$$+ \alpha(\beta_1 rw + \beta_2 txli + \beta_3 txemp + \beta_4 usemp)_{t-T-1} + \epsilon_t$$

where T is the number of lags and the term in the parentheses specifies the cointegrating relationship.

III. Data

[5]The Johansen and Juselius (1990) cointegration test was run for these two variables over the sample of 74:1 to 91:2 with the VAR lag length of 6. We did find one significant cointegrating vector at the 5 percent level. The stationary long-run relationship between relative employment and relative wage implies that the two series should not move too far apart. Any divergence from the equilibrium should be considered temporary.

The variables used to estimate the model are Texas nonfarm and U.S. nonfarm employment, TXLI, and the Texas manufacturing wage relative to the U.S. manufacturing wage. As described earlier, we construct a measure of relative wages that adjusts for the differing industrial structure in the region than in the nation. The wage variable is intended to measure the relative wage across industries that is due solely to wage differences and not to differences in industrial structure. The relative wage variable is calculated as follows:

$$RW_t = \Sigma_{i=1}^{19} \left(\frac{HWTX_{ti}}{HWUS_{ti}}\right) * \frac{EUS_{ti}}{EUS_t}$$

where RW_t is the relative wage, $HWTX_{it}$ is the hourly wage in industry i in Texas, $HWUS_{it}$ is the hourly wage in industry i in the U.S., EUS_{it} is U.S. employment in industry i, and EUS_t is U.S. manufacturing employment summed across nineteen two-digit manufacturing industries available at both the state and national levels and t is a time subscript[6]. Thus the wage variable is constructed by calculating relative wages at the two-digit manufacturing level and weighting each industry's relative wage by the industry's U.S. employment share. The wage and employment data are from the Current Employment Statistics (CES) series produced by the Bureau of Labor Statistics.

In order to compare our model to the model developed by LeSage (1990a), we calculate base employment in Texas with the following formula by LeSage:

$$RNE_{it} = RE_t \left(\frac{EUS_{it}}{EUS_t}\right)$$

where RNE_{it} is the number of employees in industry i that is necessary to supply local needs, RE_t is total regional employment, EUS_{it} is U.S. employment in industry i and EUS_t is total U.S.

[6]For Texas, wages for nineteen of the twenty two-digit manufacturing industries were available during this time period. At the national level, wages for Electric and Electronic Equipment, and Instruments and Related Products did not begin until 1988, and so altogether seventeen of the twenty industries were used.

employment. If RNE_{it} is greater than or equal to RE_{it}, nonbase employment in industry i is then equal to RE_{it} and base employment in industry i is equal to zero. If RNE_{it} is less than RE_{it}, then local employment in industry i is equal to RNE_{it} and base employment is equal to RE_{it} minus RNE_{it}.

The dynamic location quotient described above generates two time-series data which represent monthly magnitudes of local (nonbasic), denoted by LE, and export (basic) employment, denoted by XE, for the state of Texas. While LeSage used employment data in only three categories - durable, nondurable, and nonmanufacturing employment - we include employment at the two-digit SIC level in manufacturing and at the division level in non-manufacturing, for a total of twenty-nine industries. The nonfarm employment data for different sectors in Texas are from the CES. The data are seasonally adjusted with the adjustment procedure described in Berger and Phillips (1993, 1994). This adjustment controls for a break in the seasonal pattern which is often found in the state employment data.

IV. Econometric Methodology

Applying conventional VAR techniques will lead to spurious results if the variables in the system are nonstationary. The mean and variance of a nonstationary, or integrated, time series which has a stochastic trend, will depend on time. Any shock to the variable will have permanent effects on it. The most common procedure to render the series stationary is to transform it into first differences. Nevertheless, the model in first differences will be misspecified if the series are cointegrated and converge to stationary log-term equilibrium relationships.[7] The set of variables

[7]The concept of cointegration, first proposed in Granger and Weiss (1983), is fundamental to the use of the error-correction-model formulation. Engel and Granger (1987) have shown that a model estimated using differenced data will be misspecified if the variables are cointegrated and the cointegrating relationship is ignored. Cointegration means that nonstationary time series variables tend to move together such that a linear combination of them is stationary. Cointegration is sometimes interpreted as representing a long-run

in the system must cointegrate for a valid Error-Correction VAR (ECM) to exist. Therefore, tests for cointegration should be a necessary component of estimation exercises conducted with ECMs. In our analysis we test for cointegration using the general maximum likelihood approach of Johansen and Juselius (JJ) (1990).

The JJ approach differs from Engel and Granger (1987)[8], which LeSage (1990a) adopts, in that it offers an explicit criterion for choosing the number of cointegrating vectors. Suppose X_t is a (px1) vector of first difference stationary time series variables, the ECM form is

$$\Delta X_t = \Gamma_1 \Delta X_{t-1} + \dots + \Gamma_{k-1} \Delta X_{t-k+1} + \Pi_k X_{t-k} + \mu + \epsilon_t$$

where $\Gamma_i = -I + \Pi_1 + \dots + \Pi_i$ $(i=1,\dots,k-1)$ and $\Pi = \alpha\beta'$, β' is the $(p\times r)$ cointegrating vector, and α is the $(p\times r)$ vector of error correction coefficients (or speed of adjustment). Π_k is a pxp matrix and ϵ_t is a vector with p elements composed of independently and normally distributed random disturbances with zero means.

In JJ, β can be estimated and Rank(Π) can be determined by the trace test and the maximum eigen-value test. We employ both tests to check the sensitivity of the results. The former tests the null hypothesis that there are at most r cointegrating vectors against a general alternative, and the latter test the null of r cointegrating vectors against the alternative of at least $(r+1)$. Notice that if $\text{Rank}(\Pi)=r=p$, any vector is a cointegrating vector and hence all the variables in X_t are stationary. If $\text{Rank}(\Pi)=r<p$, then the series are first-difference stationary

equilibrium (steady-state) relationship.

[8]Engel and Granger (1987) proposes a cointegration test testing procedure using the Dickey-Fuller unit root test. However, their cointegration test requires prior knowledge about the cointegrating vectors, which are usually unknown. In contrast, the Johansen and Juselius (1990) multivariate testing procedure can overcome this problem. Furthermore, their method is not subject to bias due to arbitrary normalization choices since all the variables are endogenous in a VAR setting. Gonzalo (1994) concludes that Johansen and Juselius' approach has better properties than other alternative techniques after examining the asymptotic distribution of the various estimators.

and there are r cointegrating vectors. If Rank(Π) = r = 0, then no significant cointegrating vectors exist and a VAR based purely on the first difference of X_t is appropriate.

V. Empirical Results

To specify the model correctly, we check the stationarity of the log levels of all variables. Unit roots are tested for using the Augmented Dickey-Fuller method. Table 1 shows the results of both tests. Test results reveal that all variables in our model are nonstationary at the 10 percent significance level. The levels of both local employment and export employment in Texas also show strong evidence of non-stationarity. The residuals are checked for serial correlation by Q statistics. None of the Q statistics showed significant autocorrelation at the one percent level.

We apply the JJ test to the variables used in our model (rw, txli, txemp, usemp, and the variables used in the LeSage model (le, xe). Table 2 displays the results of the JJ rank tests and the critical values. In the trace test of cointegration among four variables (upper panel), the hypothesis that the number of cointegrating vectors (r) greater than 1 cannot be rejected, while the hypothesis that the number of cointegrating vectors (r) greater than or equal to 2 can be easily rejected. This finding is confirmed by the maximum eigen test which suggests rejection of the null hypothesis of r = 0 in favor of the alternative hypothesis that r = 1 at the 5 percent level. The test for the two variables in the LeSage model found one cointegrating vector, significant only at the 10 percent level. This result is similar to LeSage's study.

Table 3 presents the results of causality tests for both models. The error-correction term (EC) is added as a regressor to the VAR. To examine the serial correlation of the residuals, we apply the Ljung-Box Q tests which uniformly fail to reject the null hypothesis of no serial correlation. The lagged EC appears with a very significant coefficient in each equation. This result implies that overlooking the long-run relationship of the variables would have caused misspecification in the underlying dynamic structure. The existence of cointegration between export-based employment (xe) and local employment (le) also accords with LeSage's findings. Furthermore, the error-correction representation of Granger-causality allows for the finding, for example, that the relative wage Granger-causes Texas employment. This is true even though the coefficients on the wage variable are jointly insignificant because the coefficient of the error-correction term is significant. The effect of the relative wage on Texas employment growth, however, appears to be mostly long-run, which is consistent with our labor demand hypothesis. The Texas Leading Index has strong short-term impacts on Texas employment, which is also consistent with our hypothesis about this series. Therefore Panel A provides strong evidence of existence of bidirectional causality among the variables. In Panel B, the bidirectional causality also appears between xe and le. It is noteworthy that the evidence of causality from the export employment to the local employment is somewhat stronger statistically.

We performed several out-of-sample forecasting experiments to compare the forecasting ability of the two models. In the first experiment, we did out-of-sample forecasts at one- through thirty-six horizons. At each horizon, we updated the in-sample estimations. For instance, a twelve-step-ahead forecast of May 1992 (June 1992) was derived from the estimations using data up to May 1991 (June 1991). The one-step-ahead forecast of the same month employed the information up to April 1992 (May 1992). These iterative estimation and forecast procedures were carried out for all out-of-sample forecasts. However, to estimate the cointegrating vectors, we only use the sample for the first forecast, which covers the period from January 1974 to February 1991.

The first forecasting experiment involved the out-of-sample period from March 1991 to

March 1995.[9] One- to thirty-six-month-ahead forecasts were calculated for both models. As shown in Table 4 , the mean absolute percentage error (MAPE) for the labor-demand leading index model was smaller than that for the base model for all forecast horizons except the one-step-ahead forecast. The forecast errors for our model were generally small, with the MAPE less than 1 percent for horizons one through sixteen months. At the forecasting horizon of thirty-six months the MAPE was 1.8 percent versus 4.3 percent for the LeSage model. On average the MAPE in our model was 50 percent smaller than in the LeSage model.

A standard criticism of time series models is that they typically perform poorly at turning points. To address this issue, we looked at how each model performs in the out-of-sample periods during the last economic recession in Texas which occurred from November 1985 to March 1987. The recession coincided with a dramatic negative shock to oil prices and a tax law change which made real estate investing less profitable.[10] The diagnostics from the two models were similar to the full sample results discussed above. The variables were found to be nonstationary and cointegration was found in each of the models for the sample of January 1974 to April 1984. The results of the tests are reported in Table 1 and Table 2. Columns 3 and 4 of Table 4 show that both models performed worse during this period than in the latter period. During this period our model performed much better than the LeSage model at the short-term forecasting horizons. Changes in the leading index, which incorporated changes in oil prices and help-wanted advertising, likely helped our model to predict the quick response of the economy to the sharp decline in oil prices which began in late 1985. On average, the MAPE was 27-

[9]The employment data used in this paper are adjusted with the Berger-Phillips method as described in Berger and Phillips (1993). Thus, the data after March 1993, which are the official post-benchmarked data, are less subject to revision than they otherwise would be.

[10]For a more detailed description of the events leading to the 1986 Texas recession, see Petersen, Phillips, and Yucel (1994).

percent smaller in our model than in the LeSage model.

VI. Other Applications of the Model

While the labor-demand-leading-indicator model presented here does well in the out-of-sample forecasting experiments, the model is flexible enough to be used in other ways. For example, the model is specifically built to forecast regional employment, and thus, the forecasts of U.S. employment generated by its feedbacks with the other three variables may be less trusted. U.S. employment is likely to be forecasted better by national variables than by regional ones. However, the addition of other variables to the reduced-form VAR model may not be proper because it would cause the loss of degrees of freedom . In practice, it is easy to interject outside forecasts of U.S. employment growth into the model to potentially improve the forecast of regional employment. This is also useful if the analyst is given a U.S. forecast and assigned to making a state forecast that is consistent with it.

While this model seemed to work well for Texas, its application to other states needs further examination. As discussed earlier, many authors have found that cointegration in regional economic models is the exception rather than the rule. To address this question, we ran some preliminary tests to see if the basic long-run labor-demand model would be applicable to other states. Shoesmith (1995) carried out tests for the states of Texas, North Carolina, New York, and Vermont and found that for his five-variable model, cointegration was found only for Texas. As a comparison, we ran the basic labor-demand model shown in Equation (RW) for the other three states on monthly data from January 1974 to March 1995. Using the JJ test, we found a cointegrating relationship for all three states.[11] Furthermore, according to our findings

[11]We first tested the cointegration of the relative wage and the relative employment shown in equation (RW) for these three states. For both North Carolina and New York, cointegration was found at the 10 percent level of significance. The test result showed no significant evidence of cointegration for Vermont.

for Texas (Table 3), adding regional leading indexes to these models should help their predictive power, particularly in the short run. These results give evidence that the model presented here could successfully be applied to other states.

VII. Conclusion and Summary

Recent work in regional forecasting suggests that the ECM should be a useful tool for regional analysts. LeSage (1990a) also suggests that the ECM used within the framework of an economic-base model represents a simple, but effective, approach to regional employment forecasting. In this paper we build a simple labor-demand-leading-index ECM of Texas employment which includes national employment, the state's relative industry-weighted manufacturing wage, and the Texas Leading Index.

We find that out-of-sample errors from our model are, on average, smaller than those from an economic-base model suggested by LeSage (1990a). While the model developed here was used for Texas, similar models can be developed for most states. As noted in Phillips (1994), there are at least twenty-four state leading indexes currently being produced, and the construction of the indexes is fairly straightforward. For most states relative manufacturing wage variable can be easily constructed back to 1970 using the CES data from the Bureau of Labor Statistics. Preliminary tests suggest that the basic labor-demand model shown in equation (RW) describes a long-term cointegrating relationship in many states.

While the model presented here represents a simple and potentially effective way to forecast regional employment, there are some limitations to this approach. First, as is the case

---

However, we went further to conduct the test with an unrestricted model, namely, the coefficient for the U.S. employment was not restricted to 1. We found that cointegration existed among the three variables for all three states at the 10 percent level. Following the method described in Section III to construct the relative wage, we were able to use fifteen industries for New York, sixteen industries for North Carolina and nine industries for Vermont.

with most reduced-form time series models, the model has few policy implications. Also, we compared this model with only one other, and although we used a model that had been accepted in the literature, there is an infinite number of potential models. In general, however, the use of cyclical indicators within the context of a long-term labor-demand ECM should prove helpful to regional forecasters.

References

Berger, Frank, and Keith Phillips (1993), "Reassessing Texas Employment Growth," Federal Reserve Bank of Dallas *Southwest Economy*, July/August.

___________ and ___________ (1994), "Solving the Mystery of the Disappearing January Blip in State Employment Data," Federal reserve Bank of Dallas *Economic Review*, Second Quarter, 56-62.

Browne, Lynn E. (1989), "Shifting Regional Fortunes: The Wheel Turns," Federal Reserve Bank of Boston *New England Economic Review*, May/June, 27-40.

___________, Peter Mieszkowski, and Richard F. Syron (1980), "Regional Investment Patterns," Federal Reserve Bank of Boston *New England Economic Review*, July/August 1980.

Carlino, Gerald, and Leonard Mills (1993), "Are U.S. Regional Incomes Converging?" *Journal of Monetary Economics* 32, 335-346.

Engel, Robert, and Clive W. J. Granger (1987), "Co-integration and Error Correction: Representation, Estimation, and Testing," *Econometrica* 55, 251-76.

Fuller, W. A. (1976), *Introduction to Statistical Time Series*, New York: John Wiley.

Gonzalo, Jesus (1994), "Five Alternative Methods of Estimating Long-Run Equilibrium Relationships," *Journal of Econometrics* 60, 203-33.

Granger, C. W. J., and A. A. Weiss, (1984), "Time Series Analysis and Error Correcting Models," in S. Karlin, T. Amemiya and L. A. Goodman (eds.) *Studies in Econometrics, Time Series and Multivariate Statistics* (New York: Academic Press).

Hefner, Frank L. (1990), "Cointegration of Regional GSP with National GDP," *Studies in Economic Analysis* 13, 43-58.

Johansen, S., and K. Juselius (1990), "Maximum Likelihood Estimation and Inference on Cointegration - with Applications to the Demand for Money," *Oxford Bulletin of Economics and Statistics* 52, 169-210.

Kottman, Stacy E. (1992), "Regional Employment by Industry: Do Returns to Capital Matter?" Federal Reserve Bank of Atlanta *Economic Review*, September/October, 13-25.

Krikelas, Andrew C. (1992), "Why Regions Grow: A Review of Research on the Economic Base Model," Federal Reserve Bank of Atlanta *Economic Review*, July/August, 16-29.

LeSage, James P. (1990a), "Forecasting Metropolitan Employment Using an Export-Base Error-Correction Model," *Journal of Regional Science* 30, 307-323.

___________ (1990b), "A Comparison of the Forecasting Ability of ECM and VAR Models," *The Review of Economics and Statistics*, 664-671.

Nakosteen, Robert A., and Michael A. Zimmer (1987), "Determinants of Regional Migration by Manufacturing Firms," *Economic Inquiry*, April, 351-362.

Petersen, D'Ann M., Keith Phillips, and Mine Yucel (1994), "The Texas Construction Sector: The Tail That Wagged the Dog," Federal Reserve Bank of Dallas *Economic Review*, Second Quarter, 23-33.

Phillips, Keith (1990), "The Texas Index of Leading Economic Indicators: A Revision and Further Evaluation, Federal Reserve Bank of Dallas *Economic Review*, July, 17-25.

____________ (1994), "Regional Indexes of Leading Economic Indicators," in *Forecasting Financial and Economic Cycles*, by Michael P. Niemira and Philip A. Klein, John Wiley and Sons, Inc., 347-360.

Shoesmith, Gary L. (1992), "Noncointegration and Causality: Implications for VAR Modelling," *International Journal of Forecasting* 8, 187-199.

____________ (1995), "Long-Term Forecasting of Noncointegrated and Cointegrated Regional and National Models," *Journal of Regional Science* 35, 43-64.

Wheat, Leonard F. (1986), "The Determinants of 1963-77 Regional Manufacturing Growth: Why the South and West Grow," *Journal of Regional Science*, November, 635-659.

Table 1: Augmented Dickey-Fuller Tests

$$\Delta X_t = \alpha + (\theta t) + \rho X_{t-1} + \sum_{i=1}^{k} \beta_i \Delta X_{t-i} + \epsilon_t$$

	Trend			No Trend		
X	k	τ-statistics	Q(21)	k	τ-statistics	Q(22)
sample: January 1974–February 1991						
rw	12	-1.87	25.62	12	-2.12	26.65
txli	2	-2.23	19.10	2	-2.24	19.93
txemp	3	-1.46	29.40	3	-1.88	30.59
usemp	2	-2.33	18.14	2	-0.80	16.69
xe	3	-1.36	34.84**	3	-1.51	38.19**
le	2	-1.21	34.36**	2	-1.96	35.01**
sample: January 1974 - April 1985						
rw	12	-2.73	25.46	12	-1.48	29.64
txli	2	-1.98	14.55	2	-1.68	15.57
txemp	2	-1.28	30.43*	2	-1.05	30.84
usemp	2	-1.95	11.10	2	-0.54	10.49
xe	1	0.31	25.08	1	-1.88	25.20
le	2	-1.18	24.23	2	-0.92	24.91

Notes: All variables in natural logarithms are defined in sections 2 and 3 of the text. The choice of the lag length, k, is based on Akaike Information Criterion (AIC) in the range from 0 to 12. In all cases the null hypothesis of $\rho = 1$ can be rejected at the 10 percent significance level for the unit root tests. Ljung-Box Q statistics are reported and checked for significant autocorrelation in the residuals. The critical values for the ADF tests are presented in Fuller (1976).***, **, and * indicate significance at the 1-percent, 5-percent and 10-percent levels respectively.

Table 2: Johansen and Juselius Cointegration Tests

Chang-Phillips Model, test variables: rw, txli, txemp, usemp

Hypothesis		Statistics and Critical Values			
Null	Alt.(trace/λ-max)	λ_{trace}	$\lambda_{trace}(0.95)$	λ_{max}	$\lambda_{max}(0.95)$
r=0	r≥1/r=1	50.68**	47.18	37.98**	27.17
		52.89**		39.69**	
r=1	r≥2/r=2	12.70	25.51	5.82	20.78
		13.20		8.24	
r=2	r≥3/r=3	6.88	15.20	5.53	14.04
		4.97		4.91	
r=3	r≥4/r=4	1.35	3.96	1.35	3.96
		0.05		0.05	

Notes: The lag length in VAR is chosen to eliminate the autocorrelation of the residuals in each equation. The upper statistics are for the sample 74:1-91:2, the lower are for the sample 74:1-85:4. Variables included in the test are rw, txli, txemp, and usemp. A model that allows for linear trends is fitted to the data, which results in the cointegrating vectors without a constant term.

LeSage Model, test variables: xe, le

Hypothesis		Statistics and Critical Values			
Null	Alt.(trace/λ-max)	λ_{trace}	$\lambda_{trace}(0.90)$	λ_{max}	$\lambda_{max}(0.90)$
r=0	r≥1/r=1	18.57*	17.96	14.46*	13.78
		30.18**		28.51**	
r=1	r≥2/r=2	4.11	7.56	4.11	7.56
		1.67		1.67	

Notes: The lag length in VAR is chosen to eliminate the autocorrelation of the residuals in each equation. The upper statistics are for the sample 74:1-91:2, the lower one for the sample 74:1-85:4. Variables included in the test are xe and le.

* indicates rejection of the null hypothesis at the 10-percent level.
** indicates rejection of the null hypothesis at the 5-percent level.

Table 3: Causality Tests: monthly data 74:1-91:2
I. Panel A:

EQUATIONS

	Δrw	Δtxli	Δtxemp	Δusemp
i=1,2,3	F(3,188)	F(3,188)	F(3,188)	F(3,188)
Δrw_{t-i}	---	2.48*	0.34	1.79
$\Delta txli_{t-i}$	1.15	---	7.98***	3.35**
$\Delta txemp_{t-i}$	5.04**	0.81	---	0.54
$\Delta usemp_{t-i}$	2.25*	0.23	1.20	---
i=4	F(1,188)	F(1,188)	F(1,188)	F(1,188)
EC_{t-i}	13.15***	6.11***	6.23**	4.45**
Q(36)	35.47	36.62	38.93	22.99

II. Panel B:

	Δxe	Δle
i=1,2,3	F(3,195)	F(3,195)
Δxe_{t-i}	---	2.65**
Δle_{t-i}	3.73**	---
i=4	F(1,195)	F(1,195)
EC_{t-i}	4.56**	13.58***
Q(36)	41.04	39.18

Notes: ***, **, and * indicates significance at the 1-percent, 5-percent, and 10-percent levels respectively.

TABLE 4: OUT-OF-SAMPLE FORECASTS

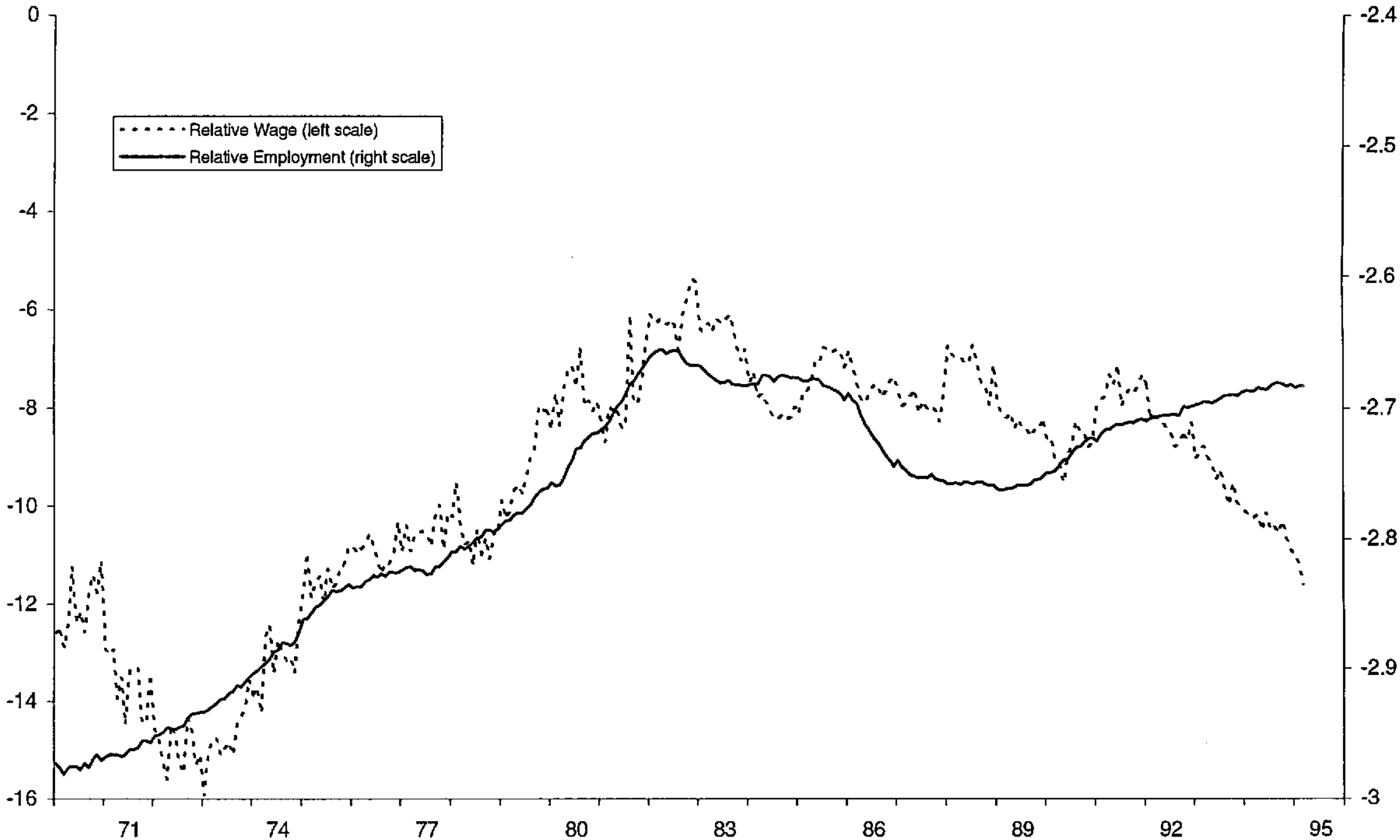
Mean Absolute Percentage Errors
Chang/Phillips and LeSage Models

Horizons	March 91 - March 95	May 85 - May 87
1	0.17(0.16)	0.25(0.30)
4	0.33(0.38)	0.74(1.38)
8	0.52(0.80)	1.83(3.45)
12	0.72(1.32)	3.32(5.55)
16	0.93(1.77)	4.68(6.95)
20	1.15(2.15)	5.30(7.42)
24	1.31(2.62)	5.83(7.58)
28	1.58(3.18)	5.83(7.04)
32	1.73(3.74)	5.65(5.77)
36	1.82(4.26)	5.66(4.97)
Average*	1.05(2.08)	4.08(5.57)

Notes:
(1) The numbers in parentheses are for the LeSage base-nonbase model. All values are in percent.
(2) The in-sample estimations for both models start from January 1974.

*The reported average is the mean of MAPEs for all thirty-six horizons.

CHART 1



Relative Wage (left scale)
Relative Employment (right scale)
0
-2
-4
-6
-8
-10
-12
-14
-16
-2.4
-2.5
-2.6
-2.7
-2.8
-2.9
-3
71
74
77
80
83
86
89
92
95

RESEARCH PAPERS OF THE RESEARCH DEPARTMENT
FEDERAL RESERVE BANK OF DALLAS

Available, at no charge, from the Research Department
Federal Reserve Bank of Dallas, P. O. Box 655906
Dallas, Texas 75265-5906

Please check the titles of the Research Papers you would like to receive:

___ 9201 Are Deep Recessions Followed by Strong Recoveries? (Mark A. Wynne and Nathan S. Balke)
___ 9202 The Case of the "Missing M2" (John V. Duca)
___ 9203 Immigrant Links to the Home Country: Implications for Trade, Welfare and Factor Rewards (David M. Gould)
___ 9204 Does Aggregate Output Have a Unit Root? (Mark A. Wynne)
___ 9205 Inflation and Its Variability: A Note (Kenneth M. Emery)
___ 9206 Budget Constrained Frontier Measures of Fiscal Equality and Efficiency in Schooling (Shawna Grosskopf, Kathy Hayes, Lori L. Taylor, William Weber)
___ 9207 The Effects of Credit Availability, Nonbank Competition, and Tax Reform on Bank Consumer Lending (John V. Duca and Bonnie Garrett)
___ 9208 On the Future Erosion of the North American Free Trade Agreement (William C. Gruben)
___ 9209 Threshold Cointegration (Nathan S. Balke and Thomas B. Fomby)
___ 9210 Cointegration and Tests of a Classical Model of Inflation in Argentina, Bolivia, Brazil, Mexico, and Peru (Raul Anibal Feliz and John H. Welch)
___ 9212 The Analysis of Fiscal Policy in Neoclassical Models[1] (Mark Wynne)
___ 9213 Measuring the Value of School Quality (Lori Taylor)
___ 9214 Forecasting Turning Points: Is a Two-State Characterization of the Business Cycle Appropriate? (Kenneth M. Emery & Evan F. Koenig)
___ 9215 Energy Security: A Comparison of Protectionist Policies (Mine K. Yücel and Carol Dahl)
___ 9216 An Analysis of the Impact of Two Fiscal Policies on the Behavior of a Dynamic Asset Market (Gregory W. Huffman)
___ 9301 Human Capital Externalities, Trade, and Economic Growth (David Gould and Roy J. Ruffin)
___ 9302 The New Face of Latin America: Financial Flows, Markets, and Institutions in the 1990s (John Welch)
___ 9303 A General Two Sector Model of Endogenous Growth with Human and Physical Capital (Eric Bond, Ping Wang, and Chong K. Yip)
___ 9304 The Political Economy of School Reform (S. Grosskopf, K. Hayes, L. Taylor, and W. Weber)
___ 9305 Money, Output, and Income Velocity (Theodore Palivos and Ping Wang)
___ 9306 Constructing an Alternative Measure of Changes in Reserve Requirement Ratios (Joseph H. Haslag and Scott E. Hein)
___ 9307 Money Demand and Relative Prices During Episodes of Hyperinflation (Ellis W. Tallman and Ping Wang)
___ 9308 On Quantity Theory Restrictions and the Signalling Value of the Money Multiplier (Joseph Haslag)
___ 9309 The Algebra of Price Stability (Nathan S. Balke and Kenneth M. Emery)
___ 9310 Does It Matter How Monetary Policy is Implemented? (Joseph H. Haslag and Scott Hein)
___ 9311 Real Effects of Money and Welfare Costs of Inflation in an Endogenously Growing Economy with Transactions Costs (Ping Wang and Chong K. Yip)
___ 9312 Borrowing Constraints, Household Debt, and Racial Discrimination in Loan Markets (John V. Duca and Stuart Rosenthal)

___ 9313 Default Risk, Dollarization, and Currency Substitution in Mexico (William Gruben and John Welch)
___ 9314 Technological Unemployment (W. Michael Cox)
___ 9315 Output, Inflation, and Stabilization in a Small Open Economy: Evidence from Mexico (John H. Rogers and Ping Wang)

___ 9316 Price Stabilization, Output Stabilization and Coordinated Monetary Policy Actions (Joseph H. Haslag)
___ 9317 An Alternative Neo-Classical Growth Model with Closed-Form Decision Rules (Gregory W. Huffman)
___ 9318 Why the Composite Index of Leading Indicators Doesn't Lead (Evan F. Koenig and Kenneth M. Emery)
___ 9319 Allocative Inefficiency and Local Government: Evidence Rejecting the Tiebout Hypothesis (Lori L. Taylor)
___ 9320 The Output Effects of Government Consumption: A Note (Mark A. Wynne)
___ 9321 Should Bond Funds be Included in M2? (John V. Duca)
___ 9322 Recessions and Recoveries in Real Business Cycle Models: Do Real Business Cycle Models Generate Cyclical Behavior? (Mark A. Wynne)
___ 9323* Retaliation, Liberalization, and Trade Wars: The Political Economy of Nonstrategic Trade Policy (David M. Gould and Graeme L. Woodbridge)
___ 9324 A General Two-Sector Model of Endogenous Growth with Human and Physical Capital: Balanced Growth and Transitional Dynamics (Eric W. Bond, Ping Wang,and Chong K. Yip)
___ 9325 Growth and Equity with Endogenous Human Capital: Taiwan's Economic Miracle Revisited (Maw-Lin Lee, Ben-Chieh Liu, and Ping Wang)
___ 9326 Clearinghouse Banks and Banknote Over-issue (Scott Freeman)
___ 9327 Coal, Natural Gas and Oil Markets after World War II: What's Old, What's New? (Mine K. Yücel and Shengyi Guo)
___ 9328 On the Optimality of Interest-Bearing Reserves in Economies of Overlapping Generations (Scott Freeman and Joseph Haslag)
___ 9329* Retaliation, Liberalization, and Trade Wars: The Political Economy of Nonstrategic Trade Policy (David M. Gould and Graeme L. Woodbridge) (Reprint of 9323 in error)
___ 9330 On the Existence of Nonoptimal Equilibria in Dynamic Stochastic Economies (Jeremy Greenwood and Gregory W. Huffman)
___ 9331 The Credibility and Performance of Unilateral Target Zones: A Comparison of the Mexican and Chilean Cases (Raul A. Feliz and John H. Welch)
___ 9332 Endogenous Growth and International Trade (Roy J. Ruffin)
___ 9333 Wealth Effects, Heterogeneity and Dynamic Fiscal Policy (Zsolt Becsi)
___ 9334 The Inefficiency of Seigniorage from Required Reserves (Scott Freeman)
___ 9335 Problems of Testing Fiscal Solvency in High Inflation Economies: Evidence from Argentina, Brazil, and Mexico (John H. Welch)
___ 9336 Income Taxes as Reciprocal Tariffs (W. Michael Cox, David M. Gould, and Roy J. Ruffin)
___ 9337 Assessing the Economic Cost of Unilateral Oil Conservation (Stephen P.A. Brown and Hillard G. Huntington)
___ 9338 Exchange Rate Uncertainty and Economic Growth in Latin America (Darryl McLeod and John H. Welch)
___ 9339 Searching for a Stable M2-Demand Equation (Evan F. Koenig)
___ 9340 A Survey of Measurement Biases in Price Indexes (Mark A. Wynne and Fiona Sigalla)
___ 9341 Are Net Discount Rates Stationary?: Some Further Evidence (Joseph H. Haslag, Michael Nieswiadomy, and D. J. Slottje)
___ 9342 On the Fluctuations Induced by Majority Voting (Gregory W. Huffman)
___ 9401 Adding Bond Funds to M2 in the P-Star Model of Inflation (Zsolt Becsi and John Duca)
___ 9402 Capacity Utilization and the Evolution of Manufacturing Output: A Closer Look at the "Bounce-Back Effect" (Evan F. Koenig)
___ 9403 The Disappearing January Blip and Other State Employment Mysteries (Frank Berger and Keith R. Phillips)
___ 9404 Energy Policy: Does it Achieve its Intended Goals? (Mine Yücel and Shengyi Guo)
___ 9405 Protecting Social Interest in Free Invention (Stephen P.A. Brown and William C. Gruben)
___ 9406 The Dynamics of Recoveries (Nathan S. Balke and Mark A. Wynne)
___ 9407 Fiscal Policy in More General Equilibriium (Jim Dolman and Mark Wynne)
___ 9408 On the Political Economy of School Deregulation (Shawna Grosskopf, Kathy Hayes, Lori Taylor, and William Weber)
___ 9409 The Role of Intellectual Property Rights in Economic Growth (David M. Gould and William C. Gruben)
___ 9410 U.S. Banks, Competition, and the Mexican Banking System: How Much Will NAFTA Matter? (William C. Gruben, John H. Welch and Jeffery W. Gunther)
___ 9411 Monetary Base Rules: The Currency Caveat (R. W. Hafer, Joseph H. Haslag, andScott E. Hein)
___ 9412 The Information Content of the Paper-Bill Spread (Kenneth M. Emery)

___ 9413 The Role of Tax Policy in the Boom/Bust Cycle of the Texas Construction Sector (D'Ann Petersen, Keith Phillips and Mine Yücel)
___ 9414 The P* Model of Inflation, Revisited (Evan F. Koenig)
___ 9415 The Effects of Monetary Policy in a Model with Reserve Requirements (Joseph H. Haslag)
___ 9501 An Equilibrium Analysis of Central Bank Independence and Inflation (Gregory W. Huffman)
___ 9502 Inflation and Intermediation in a Model with Endogenous Growth (Joseph H. Haslag)
___ 9503 Country-Bashing Tariffs: Do Bilateral Trade Deficits Matter? (W. Michael Cox and Roy J. Ruffin)
___ 9504 Building a Regional Forecasting Model Utilizing Long-Term Relationships and Short-Term Indicators (Keith R. Phillips and Chih-Ping Chang)

Name:	**Organization:**
Address:	**City, State and Zip Code:**
Please add me to your mailing list to receive future Research Papers: _____ **Yes** _____ **No**	

Research Papers Presented at the
1994 Texas Conference on Monetary Economics
April 23-24, 1994
held at the Federal Reserve Bank of Dallas, Dallas, Texas

Available, at no charge, from the Research Department
Federal Reserve Bank of Dallas, P. O. Box 655906
Dallas, Texas 75265-5906

Please check the titles of the Research Papers you would like to receive:

____ 1 A Sticky-Price Manifesto (Laurence Ball and N. Gregory Mankiw)

____ 2 Sequential Markets and the Suboptimality of the Friedman Rule (Stephen D. Williamson)

____ 3 Sources of Real Exchange Rate Fluctuations: How Important Are Nominal Shocks? (Richard Clarida and Jordi Gali)

____ 4 On Leading Indicators: Getting It Straight (Mark A. Thoma and Jo Anna Gray)

____ 5 The Effects of Monetary Policy Shocks: Evidence From the Flow of Funds (Lawrence J. Christiano, Martin Eichenbaum and Charles Evans)

Name:	Organization:
Address:	City, State and Zip Code:
Please add me to your mailing list to receive future Research Papers:	______Yes ______ No