The Long Run Shift-Share: Modeling the Sources of Metropolitan Sectoral Fluctuations.

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ABSTRACT: Attempts to analyze the sources of regional-sectoral fluctuations have been reductions of a VAR that have simplified either the VAR's orthogonalization or long-run multiplier matrix. These *shift-share decompositions* are shown to be largely rejected by the data, mainly because they fail to properly account for local supply shocks. However the violence done to the data is more severe in short run than in the long, making the long run shift share a viable structure.

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1. Introduction

In this tribute to the career of Robert Engle, attention should be given to an aspect of his early career that is not universally recognized, that of urban and regional economist. As related in his interview with Diebold (2003), upon arriving at MIT Engle was asked by Franklin Fisher and Jerome Rothenberg to collaborate on the construction of a multi-equation structural model of the Massachusetts economy, and this led to a number of publications at the outset of his career. His involvement with this subfield did not end there. An examination of his curriculum vitae reveals that of his first thirteen publications, seven were in the field of urban economics, and there are many more publications in that area through the early 1990s. Perhaps of equal interest is the fact that many of his contributions to "pure" econometrics used urban and regional data to illustrate the methods associated with those contributions. Two prominent examples are his paper on the parameter variation across the frequency domain (Engle 1978) and Engle and Watson (1981) which introduced the DYMIMIC model. As he notes in the interview with Diebold, "there is wonderful data in urban economics that provides a great place for econometric analysis. In urban economics we have time series by local areas,and wonderful cross sections...".

One of the natural links between urban economics and time series econometrics is the examination of urban fluctuations. Because of data availability, such analysis focuses on the determination of metropolitan area employment and labor earnings, and, again because of the data, sectoral level data is often employed in the analysis. This is helpful and appropriate, because both journalistic and academic explanations of the differences in cyclical movements of aggregate urban employment often center on differences in sectoral composition across metropolitan areas. On that account the focus turns to the sources fluctuations in metropolitan industry sectors. For eample Brown, Coulson and Engle (1991), following Brown (1986), ask the basic question of whether or not metropolitan industry sectors are cointegrated (Engle and Granger, 1987) with national industry counterparts, and Brown, Coulson and Engle (1992) ask, in part, under what circumstances metropolitan sectoral employment is cointegrated with aggregate metropolitan employment.

In what follows, methods are proposed to delineate the sources of sectoral fluctuations in metropolitan economies. This delineation has four steps. First, a general VAR is constructed which accounts for both short and long run fluctuations. Second, a large number of "traditional" models of regional economics (including the two cointegration problems of the preceding paragraph) are shown to be reductions of this general VAR, although a by-product of the analysis is that it is not likely that all of these reductions can be applied simultaneously. Both of these steps occur in the next section. In Section 3 the restrictions implied by the restrictions of the traditional model are tested using data from 10 sectors and five cities. None is found to be universally applicable, though some do less violence to the data than others. Given these results, the fourth step of estimating the complete VAR (for each city industry) is undertaken under five different assumptions. The overall result is that the traditional models are unsatisfactory because the neglect the role of local supply shocks, although this neglect does more damage in "short run" models than in those that invoke cointegration.

2. A general model and some specializations

The goal of this analysis is to estimate the sources of sectoral fluctuations in a metropolitan area– for example, the Los Angeles manufacturing sector. Such sources can be conveniently catalogued as arising from four different levels: national (aggregate US), industrial (US manufacturing), metropolitan (aggregate Los Angeles), and sources that are idiosyncratic to the particular metropolitan sector. Consider, then, the following vector autoregression (VAR), which for simplicity is restricted to first order autoregressive processes (an assumption relaxed in its empirical implementation):

$$\begin{pmatrix} \Delta n \\ \Delta i \\ \Delta m \\ \Delta e \end{pmatrix}_t = \begin{pmatrix} k_1 \\ k_2 \\ k_3 \\ k_4 \end{pmatrix} + A_1 \begin{pmatrix} \Delta n \\ \Delta i \\ \Delta m \\ \Delta e \end{pmatrix}_{t-1} + A_0 \begin{pmatrix} n \\ i \\ m \\ e \end{pmatrix}_{t-1} + \begin{pmatrix} u_n \\ u_i \\ u_m \\ u_e \end{pmatrix}_t$$

$$\operatorname{cov} \begin{pmatrix} u_n \\ u_i \\ u_m \\ u_e \end{pmatrix} = \Omega \qquad (1)$$

where

n_t = log of aggregate national employment at time t

i_t = log of national employment in the same specified industry at time t

m_t = log of aggregate local employment at time t

e_t = log of local employment in a specified industry at time t,

and the k_i are intercept terms.

Consideration of this issue has been one of the primary concerns of empirical regional economics over the past half century[1]. In that period of time, a number of models have been developed, that in essence impose extra structure on the parameters of (1). In the extreme, such simplifications become *shift-share* decompositions, requiring no estimation at all. The exact form of the shift-share decomposition varies somewhat. Our baseline version of this decomposition seems to originate in Emmerson, Ramanathan and Ramm (1975):

$$\Delta e_t = \Delta n_t + (\Delta i_t - \Delta n_t) + (\Delta m_t - \Delta n_t) + (\Delta e_t - \Delta m_t - \Delta i_t + \Delta n_t) \tag{2}$$

Growth in a local industry is decomposed into four parts. The first component, the national component, estimates the impact of national employment movements on local employment movements. If, say, national employment grows at 5% in a year, then, other things equal, the local industry– say the finance sector in Boston-- is also expected to grow at the same 5% rate. The second component, the industry component, is the deviation of the national industry growth rate from that of the nation as a whole. Thus if the national finance sector grew at a rate of 10%, then the Boston finance sector should, other things equal, also be expected to grow at that same rate, with national and industry factors each responsible for half of that growth. Similarly, the third component is dubbed by Dunn (1960) the total share component, and is the deviation of the overall metropolitan growth rate from the national growth rate; obviously this is the contribution of local

[1]It should be noted at the outset that such a model can only be used to assess the sources of fluctuations of e_t, and not the other three series, all of which include e_t in their aggregations. A finding that e was a source of fluctuations of n, m, or i would seem to be vacuous without consideration of the impact of other industries or locations. For an analysis of the reverse question, industry and regional impacts on national fluctuations, see e.g. Horvath and Verbrugge, Altonji and Ham, Norrbin and Schlagelhauf. At the metropolitan level the role of sectoral fluctuations in determining aggregate metropolitan employment see Coulson (1999) and Carlino, DeFina and Sill (2000)

aggregate growth to local sector growth. The fourth component is the change in the industry's share of employment at the metropolitan level relative to its share at the national level. It is the percentage change in the familiar location quotient and is interpretable as the outcome of local supply shocks to local employment growth (given that the total share components nets out local demand factors and the industry component presumably nets out technology shocks that are common to all locations).

How can the shift-share model be used to inform the specification of the VAR (1)? There are effectively two general approaches which are not mutually exclusive, though for purposes of this paper they will be. One is to view (2) as an *orthogonalization;* that is, each of the components is assumed to be uncorrelated with the others, and therefore capable of separate study. How this has happened in the historical literature will be addressed below, but for the moment note that in the context of the VAR, the implications of this (Coulson (1993)) is that we should premultiply both sides of (1) by the orthogonalization matrix W, where

$$W = \begin{pmatrix} 1 & 0 & 0 & 0 \\ -1 & 1 & 0 & 0 \\ -1 & 0 & 1 & 0 \\ 1 & -1 & -1 & 1 \end{pmatrix} \qquad (3)$$

and we have

$$W\begin{pmatrix}\Delta n\\ \Delta i\\ \Delta m\\ \Delta e\end{pmatrix}_t = Wk + WA_1\begin{pmatrix}\Delta n\\ \Delta i\\ \Delta m\\ \Delta e\end{pmatrix}_{t-1} + WA_0\begin{pmatrix}n\\ i\\ m\\ e\end{pmatrix}_{t-1} + \begin{pmatrix}e_n\\ e_i\\ e_m\\ e_e\end{pmatrix}_t \qquad (4)$$

where k is the vector representation of the intercept terms,

$$u=W^{-1}e \qquad (5)$$

and the components of e are orthogonal. Thus we can write

$$\Omega = W^{-1}DW'^{-1} \qquad (6)$$

The orthogonalization of the VAR is much the same that occurs in ordinary VARs in that the orthogonalization matrix is triangular, however, given the nature of the homogeneity restrictions, the model is an overidentified structural (B-form) VAR (Coulson, 1993, Lutkepohl, 2003). The reasonableness of the structure, which is equivalent to testing the overidentifying restrictions, is also a test of the reasonableness of separately analyzing the components of the shift share decomposition, as is typically the case, even today.

As it happens, models and modes of regional analysis that view shift-share through this lens very often make implicit (and sometimes explicit) assumptions on the nature of the long run behavior of the components, that is to say, on the form of the matrix A_0. As is well-known, the rank of A_0 is critical to the time series representation of the vector of variables. If this rank is zero the variables are all integrated of order 1 (at least) and are not cointegrated; it happens that this is the explicit assumption of many previous models, as noted in Brown, Coulson and Engle (1991). It is for this reason that shift-share is regarded as a short-run model. Long run considerations, as manifested in A_0, are non-existent

If the rank is positive but less than the full rank of four, there is cointegration among the variables. If the rank is full, then the variables are ostensibly stationary. It will be demonstrated later that this last possibility will not trouble us much, and so if $A_0 \neq 0$ then the proper question is how many cointegrating vectors exist within the system. Let the four components of the data vector be notated as x_t. The essence of cointegration is that the number of variables in x is greater than the number of integrated processes characterizing their evolution, therefore the components of x are tied together in the long run. This is delivered by the fact that while each of the x variables is I(1), certain linear combinations are stationary, that is I(0). If those combinations are notated as $\beta' x$ we can write

$$A_0 = \alpha\beta' \tag{7}$$

where β is the kxr matrix of the r *cointegrating vectors*, and α is the kxr matrix of *adjustment speeds*.[2] As is well known, α and β are not separately identified (since for any non-singular 4x4 matrix F, the two vectors $\alpha^*=\alpha F$ and $\beta^*= F^{-1}\beta$ would be observationally equivalent to α and β.

The usual procedure is to specifiy restrictions on β, which are usually zero or normalization restrictions. To anticipate the implications of the long-run shift share model, we suppose that in our system of four variables we have three cointegrating vectors. A_0 would therefore have rank=3 and 16-1=15 free parameters. The matrix of adjustment speeds, α, is typically freely estimated, and therefore uses up 4x3=12 parameters, leaving β with three. Typically, then, the cointegrating vectors would be given, without loss of generality, as

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[2]Note that we could write the levels term as $\alpha(\beta' x_t)$. The parenthetic part is known as the error correction term, and is a measure of the distance of the x vector from its long run equilibrium. The α term is then described as the *speed of adjustment* and it, as the name suggests, is a measure of how fast the error correction term goes to its equilibrium value of zero.

$$\beta = \begin{pmatrix} \beta_1 & 1 & 0 & 0 \\ \beta_2 & 0 & 1 & 0 \\ \beta_3 & 0 & 0 & 1 \end{pmatrix} \tag{8}$$

This is, of course, where the shift-share decomposition comes in. The second strand of models that deal with shift-share analysis have used the decomposition to identify, and over-identify, the matrix β. Accumulate and slightly rearrange the decomposition (2) to obtain the identity:

$$(e_t - n_t) = (i_t - n_t) + (m_t - n_t) + (e_t - m_t - i_t + n_t) \tag{9}$$

The idea is that now each of the parenthetical terms represents a cointegrating vector; that while each of the data series is I(1), the differences displayed represent stationary objects. Equally obvious is the fact that if any three of the parenthetic terms are stationary the fourth one is as well, and so one can, indeed must, be omitted from the rows of β. In the standard formulation (8), this *long run shift share* model would impose the further restrictions $\beta_1 = \beta_2 = \beta_3 = -1$. But clearly we could implement the alternative formulation:

$$\beta = \begin{pmatrix} -1 & 1 & 0 & 0 \\ -1 & 0 & 1 & 0 \\ 1 & -1 & -1 & 1 \end{pmatrix} \tag{10}$$

which implies that the industry component, the total share and the location quotient are all stationary. This form of β is attractive in that it is simply the last three rows of W.

It should be noted that the existence of three cointegrating regressions in four variables implies that the entire system is driven in the long run by one shock. Given the implicit assumptions on causality that are inherent in the W matrix, that shock is the one to national employment. This seems somewhat implausible, so (as in the short run model) we can consider other parameterizations to this model as alternatives to the long run shift share, models that assume some cointegration, but not "full cointegration" as implied by the long run shift share model.

To summarize: the short run shift share model implies (a) that the rank of A_0=0 so that the model is one of changes; and (b) an orthogonalization of those data series that involves homogeneity restrictions. The long-run shift share implies (a) the rank of A_0=3, and (b) similar homogeneity restrictions on the cointegrating matrix β.

We can now survey the historical development of the shift-share model as a series of restrictions on the above delineated types. It should not be assumed that the authors who are cited as developing these various models necessarily found evidence in their favor, only that they developed and used them for testing or forecasting purposes.

1. Dunn (1960): The total share model

Dunn (1960) views the shift-share model as a model of total regional employment rather than local sectoral employment. He proposes the following decomposition[3]:

[3]Actually, Dunn (1960) and much of the literature that follows frames shift and share in terms of numbers of jobs gained or lost. Thus they would premultiply both sides of (2) by m_{t-1} (or later by e_{t-1}) . In the interest of simplicity this modification is ignored.

$$\Delta m = \Delta n + (\Delta m - \Delta n) \tag{11}$$

With m/n as the *share* of the region in the national economy, the second term is, naturally enough, the *shift* in that share. Hence the name. Because this needs to be distinguished from industry based shift-share, these are actually dubbed the total shift and share. Given the language in Dunn (1960) this model is viewed as one in which, other things equal, the region should grow at the same rate as the nation as a whole. Dunn clearly views the model as one of the short run, hence we would view the decomposition as a reduction of the orthogonalization scheme above, specifically W_{31}=-1 and W_{32}=0.

The total share model does not operate at the level of the industry (either local or national). This is not at all the same thing as assuming that industry effects are non-existent, merely that they are not part of the assumed structure. Thus the W matrix is written as:

$$W = \begin{pmatrix} 1 & 0 & 0 & 0 \\ \beta_{21} & 1 & 0 & 0 \\ -1 & 0 & 1 & 0 \\ \beta_{41} & \beta_{42} & \beta_{43} & 1 \end{pmatrix}$$

Also, there is no cointegration between m and n, and thus the total share is non-stationary. In the first order model above this implies that the share is a random walk.

2. Carlino and Mills (1993): Long run constant total share (stochastic convergence)

In direct contrast to Dunn's implicit assumption that the total share is a random walk, Carlino and Mills (1993) test for the proposition that the total share m/n is stationary[4]. Thus the share held by a particular region is constant in the long run. This is taken as evidence of stochastic convergence, the idea being that deviations from this share will not persist. The long run constant total share model is therefore manifested as a restriction on the rank of β=1, as there is only one long run restriction, and this row will be of the form (-1 0 1 0); that is, neither e_t nor i_t is expected to be part of the long run model.

3. H. Brown (1969): Sectoral shift-share

Brown (1969) introduced the three part shift-share model, which shifted focus from the total regional share to the industry share:

$$\Delta e = \Delta n + (\Delta i - \Delta n) + (\Delta e - \Delta i) \tag{12}$$

with attention focusing on the behavior of the final term, the *regional shift*, which is easily seen to be

[4]Though perhaps with a structural break.

the change in the industry share (e/i) held by the region. The fact that these three terms were regarded as separately analyzed series is an implicit assumption that the decomposition is in fact an orthogonal one (Coulson 1993). Noting that m plays no role in this decomposition the W^{-1} matrix is of the form

$$W^{-1} = \begin{pmatrix} 1 & 0 & 0 & 0 \\ -1 & 1 & 0 & 0 \\ w_{31} & w_{32} & 1 & 0 \\ 0 & -1 & 0 & 1 \end{pmatrix} \qquad (13)$$

Once the three part decomposition is developed, the assumption of orthogonality becomes explicit, as modeling of the shift component e/i is now the focus of the research program. Not only that, but the short run assumption also becomes operational. In an attempt to frame shift-share as a forecasting tool, Brown (1969) postulated that the region's industry share was a random walk. This implies not only the orthogonalization suggested in (12) but that there is no cointegration between e and i.

4. S. Brown, Coulson and Engle (1991): Constant share

This is the natural counterpart to the martingale share model, implying that the orthogonalization in (12) is appropriate at the long run rather than short run time horizon, that is that there is a single cointegrating vector in β, and that the row is of the form (0 1 0 -1). Brown, Coulson and Engle (1991) frame this model as a set of regional production functions with a fixed-in-location factor of production, Technology shocks are national, and migration across regions equilibrates the share of production for each region in the long run as a function of the region's share in the fixed factor.

5. Sasaki (1963): The base multiplier model

One of the workhorse models of regional economics is the base multiplier model, which implies a relationship between (a set of) local industry employments (the basic sectors) and aggregate regional employment. This is not a relationship between e and m, per se, but between the sum of a subset of e's and m and was first placed in a regression context by Sasaki (1963). Nevertheless, while a regression of m on e would yield a biased estimate of the multiplier (as the intercept term) if the base multiplier theory holds there should still be a unit elastic relationship between each sectoral employment and total regional employment in the short run.

6. Brown, Coulson and Engle (1992): Long run base multiplier

If the employment series are integrated, then Brown, Coulson and Engle (1992) demonstrate that the base multiplier model implies that e and m will be cointegrated, regardless of whether e is not part of the basic sectors, if certain other conditions hold, to be discussed shortly. Thus there will be a row of β that can be rendered as (0 1 0 -1).

It is of interest to note that the combination of restrictions implied in models 2, 4 and 6 yield the long-run shift share model. The three rows of β discussed in those models are linearly independent and are equivalent to the matrix in equation (10). As a further interesting demonstration of this note that

$$(e_t - m_t) = (n_t - m_t) + (e_t - i_t) + (i_t - n_t)$$

Thus the three models together imply that national industries are cointegrated with the national aggregate. These seems implausible on its face, since it would imply that technology shocks are identical across industries.

Thus, one of the three long run models must be wrong. Our preliminary candidate is model (6), the long run base multiplier. The "certain other conditions" alluded to above have to do with the cointegration of the various sectors at the local level. Basically, if e is a basic sector, then it must be cointegrated with other basic sectors, again implying something like common technology shocks. If e is a local sector, it must be cointegrated with other local sectors, which presumably implies common demand shocks. At the local level this is slightly more plausible; nevertheless, the long-run shift share model does require a lot of the data.

To round out the model descriptions we reiterate the full models described in the beginning:

7. Coulson: The four part, nonstationary shift- share model

The three part decomposition/orthogonalization (model 5) is unsuitable particularly because it is difficult to interpret the components of the decomposition in a coherent manner. If e and i are employments in an export-oriented industry then a change in the share plausibly represent supply shocks to that region-industry (at least relative to the national industry); but if there are regional demand shocks for the local output, then the shift term will conflate them with the supply shocks. As noted, the four part decomposition originated by Emmerson, Ramanathan and Ramm (1975) overcomes this problem by re-introducing the total shift and re-orthogonalizing (as particularly noted by Berzeg (1978)) with W is as described in (3) above. Thus this four-part model is basically a pair of hypotheses: (a) that there is no cointegration among the four variables in levels, and thus that the VAR should be specified in changes; (b) that the matrix W describes the appropriate orthogonalization.

8. Long run shift share:

The long run counterpart to Coulson (1993) is the long run shift-share, as previously discussed[5]. There are three maintained hypothesis, that (a) the data is integrated; (b) there are three cointegrating vectors among the four variables; (c) that (10) describes the cointegrating relationships.

3. Data and Evidence.

Data on full-time employment is drawn from the website of the US Bureau of Labor Statistics (www.bls.gov). MSA data is drawn from five different MSAs: Philadelphia, Dallas, Atlanta, Chicago, and Los Angeles. Not every industry is available in every MSA, so for purposes of comparability,

[5]There are several other variants on the above models, but are omitted from the present survey. Beaud (1964) argues that the shift itself is a random walk, which would indicate that the employment series are I(2). Test results (not included) indicate no evidence of this. Theil and Ghosh (1980) model the decomposition in effect as a two-way anova model, where the interaction term, i.e. the location quotient, plays no role.

we use the broad industry aggregates ("supersectors") of the NAICS, which are listed in the Tables. Comparable data is drawn from the US employment page for aggregate and supersector employment. The data is monthly and ranges from January 1990 through August of 2006.

Our first task is to determine the integratedness of the series in question. All of the models above implicitly assume that the series are indeed integrated. Table 1 presents augmented Dickey-Fuller tests for each of the series. The Dickey-Fuller test is a test for stationarity, regressing the change in a variable on the lagged level (i.e. a univariate version of the final equation in the VAR(1)). Rejection of the null hypothesis indicates that the series in question is stationary. As can be seen, the test-statistics are almost invariably below the 5% critical value. Of the 72 series listed in Table 1, four have test-values greater than the 5% critical value, about what would be expected if the null were universally true and the tests were independent (which of course they are not). The general conclusion is therefore that the series are indeed integrated.

This paves the way for Table 2, which tests the extent to which the four series are cointegrated with each other. The unrestricted VAR (1) is estimated using each city-industry and the three more aggregated series that correspond to it[6]. Trace tests (Johansen, 1985) are then performed sequentially to reject or fail to reject whether the rank of the matrix A_0 is zero, one, or two. That is, zero rank is tested, and if rejected, a rank of one is tested, and if that is rejected a rank of two is tested. Given that a rank of four is only possible if the data is stationary, testing ceases if a rank of

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[6]Equation (1) contains intercept terms. The VARs are estimated under the assumption that part of this intercept is "inside" the cointegrating relation and part is "outside" (which are not separately identified). The first part then corresponds (under the homogeneity assumption at least) the proportionalities which exist across different levels of aggregation, and the second is congruent with the assumption that the employment levels have deterministic trends. The models are estimated using Eviews, which identifies the first part by assuming that the "inside" part is zero during estimation, and then regressing the error correction term on an intercept. The difference between that and the estimated constant becomes the trend term.

two is rejected in favor of a rank of three. Recall that the long run shift share hypothesis is that the rank of the matrix is three.

Five points can be made from Table 2:

1. There is cointegration of some kind in almost all of the VARs. Only one combination, that associated with Atlanta's Trade, Transport and Utilities sector, failed to reject the null hypothesis that the rank of A_0 was zero, and the prob-value of that test was 6.5%

2. At the other extreme, there is only a modest amount of empirical support for the long-run shift-share model, in that relatively few of the VARs exhibit three cointegrating vectors. This is to be expected given the discussion above.

3. Nevertheless, there are patterns to the number of cointegrating vectors. More cointegration (and more evidence of the long-run shift share model) are observable in the Construction and Government sectors and in the Los Angeles MSA. Other industries (infomratoin services, finance, other services) and cities exhibit much less cointegration.

4. A question of importance is the extent to which the results from point 3 are influenced by the results from Table 1. For instance, Los Angeles has more cointegration than other cities, but it is also one of the two cities where the unit root null was rejected for its aggregate employment. On the other hand, Dallas aggregate employment also was not found to have a unit root, and its VARs exhibit considerably less cointegration than does Los Angeles. Similarly, aggregate employment in the US finance sector also appeared to be stationary, and yet across the MSAs, the finance sectors VARs exhibit much less cointegration than the construction sector, or indeed any sector. The

bottom line is that very little about Table 3 could have been inferred a priori from the results in Table 1.

5. Since the predominant finding is that these VARs exhibit one or two cointegrating relationships, it would be prudent to use the bivariate cointegrating models 2, 4, and 6 to seek a better understanding. Tables 3 and 4 pursue this course.

The first row of table 3 provides trace test-statistics for stochastic convergence of the indicated MSAs, i.e of long run constant total share, vs. the non-stationary version of Dunn (1960). An asterisk indicates rejection at the 5% level of the null hypothesis that there is no cointegration between aggregate city employment and aggregate US employment. As can be seen, this is the case for three of the five cities (Philadelphia, Atlanta and Los Angeles). In addition for those cases where the null is rejected, we further test whether the un-normalized coefficient in the cointegrating regression is equal to one. This is performed by asking if one is in the 95% confidence interval of the regression coefficient. If so, a notation of =1 is made, and ≠1 if not. As the Table demonstrates, this is true of Los Angeles but not of the other two cities. The remaining rows of Table 3 are tests of long run constant industry share.(as in Brown, Coulson and Engle (1991). Again, asterisks indicate cointegration, and the test of unity are indicated as well.

What conclusions can be drawn from Table 3?

The industry level results have a broad range of interpretations. At one extreme is the Education and Health Services sector, in which all five city employments are cointegrated with national employment, and four of those are statistically indistinguishable from the constant share model. An interpretation of this is of course that permanent shocks, i.e. productivity shocks, occur at the national level and percolate immediately to each local industry, and local productivity shocks are unimportant. At the other extreme is the information sector, where no local sector is cointegrated with the aggregate. An interpretation of this result is that productivity shocks are completely local; there is no national trend to tie local sectors to the broader. While a few industries display results that are to an extent like those of the information sector (e.g. nondurable manufacturing), the most common outcome is a mixture of noncointegration and cointegration with a nonunit coefficient. For example, Professional and Business Services exhibits five cities with no cointegration and three with nonunit cointegration. Aside from the difficulties of interpreting nonunit cointegration (as a partial adoption of national technology shocks?) the variety of responses makes it supremely difficult to draw general conclusions.

At the aggregate level we see that only Los Angeles fails to reject the homogeneity requirement for the constant long run share model, while Dallas and Chicago, at the other extreme, are not cointegrated at all with national employment. Again the absence of similar results across cities makes generalities impossible. But even so, puzzles arise. For example, Los Angeles conforms to the long run total share model model even though none of the component industries do.

Table 4 provides tests of the Brown, Coulson, Engle model of long run base multipliers. There is very little evidence of cointegration (aside from Los Angeles, and the construction and business service sectors) and almost no evidence of unit responses (only two cases). This, as noted, is to be expected. Under the supposition that permanent components of employment series are due to productivity shocks, it is quite natural for there to be cointegration between local and national sectors in the same industry; it would be quite another for different industries in the same city to have such a correspondence. As Brown, Coulson and Engle (1992) note, it is possible for a single local industry series to be cointegrated with its metropolitan aggregate. The example discussed there concerned a single basic sector, which could be cointegrated with metropolitan employment if it was cointegrated with the other basic sectors. Such a scenario, as noted is quite unlikely, since the productivity shocks are unlikely to be the same across industries. What is perhaps more likely is a second example (only indirectly discussed in Brown, Coulson and Engle (1992)) where a single local sector can be cointegrated with the aggregate if it is cointegrated with other local-serving industries. This is more plausible only in the sense that local-serving industries are largely in the service sector, and the permanent shocks that accrue to them are perhaps more likely to be demand shocks. By this reasoning it is perhaps sensible that the cointegration that does occur is in two sectors that are plausibly local-serving: construction and business services.

Obviously, neither the long run shift share nor the short run shift share fully describe the fluctuations of regional economies. In order to say more, the VAR itself must actually be estimated. We perform five VARs for each city industry VAR.

(A) The short run shift share: the model is estimated in differences, and the orthogonalization (3) is imposed.

(B) The short run VAR: the model is estimated in differences and only a causal ordering implied by W is imposed (i.e. without the homogeneity restrictions).

(C) The intermediate model: Cointegration is assumed, with the number of cointegrating vectors as indicated by Table 2. Statistically, this is the "correct" model.

(D) The long run shift share: Three cointegrating relations are assumed and the homogeneity restrictions are added[7].

The VARs are estimated using 4 lags as before, and compared using the 24-month forecast error variance decomposition. The results for the six sampled MSAs are in Tables 5 through 9. The variation of results across cities and industries is large. The following stylized conclusions might be drawn, although every one of these has exceptions.

A good starting point is the comparison of Models A and B, both of which assume a lack of cointegration, but which differ in whether they impose the shift share orthogonalization (A) or not (B). Not shown in the tables is that the overidentifying restrictions that impose the normalization are universally rejected. The shift-share model is not appropriate; in comparing models A and B we see that the statistically preferred model B assigns far more explanatory power, on average, to the local industry shock, and (less regularly, the aggregate metro shock) than model A. This is natural; what the shift share model does is force local movements to follow movements in broader aggregates in a one-for-one manner, thus ignoring the role of local supply shocks. This can be contradictory to the actual movements of local industries, and thus not imposing the short run constraints would seem to be preferable. Another way of looking at this is to note that in the first step of the variance decomposition in Model A, all four shocks are given equal weight (as per the structure of the matrix W), and the force of this persists even to the 24-step horizon.

When the statistically preferred number of cointegrating vectors are assumed to exist (Model C), the results are generally closer to the results in Model B than to those in Model A. Generally, though, Model C does assign more explanatory power to national and national-industry shocks than does B. This is to be expected given the previous bivariate results of Tables 3 and 4. Note that

[7]A fifth model was estimated, which provided for three cointegrating relations, but without imposing the homogeneity restriction of the long run shift-share. As might be expected, the results were intermediate between those of Models C and D.

bivariate cointegration was far more common in the relationship between local industry and national industry than between local industry and the aggregate local economy. Thus we would expect that when cointegration is allowed into the system, that the impact of the nation and national industry would increase. By and large (but by no means universally) this result is confirmed.

As we move from model C to model D, recall that two modeling changes are made. First the number of cointegrating vectors is forced to be three. This would not be expected to make much of a difference in the results, since the extra cointegrating coefficient would presumably be close to zero. The imposition of unit coefficients (especially when they would otherwise be zero) is therefore presumably of more importance. Note first of all (test statistics not shown) that these unitary restrictions are universally rejected by the data at any conventional level of significance. Second, while there are strong differences in the results, these results do not appear to have any systematic pattern. In particular the share of the forecast error variance that is absorbed by the idiosyncratic shock does not show systematic rise or fall when the long run shift-share restrictions are imposed. Thus the imposition of the long run shift share might be particularly dangerous, since there is little indication of which direction the bias from the model runs.

4. Summary and Concludions

A natural intersection of urban economics and time series econometrics is in the examination of urban fluctuations. In this paper, the work of Robert Engle at this intersection is carried forward. The traditional models of metropolitan sectoral fluctuations investigated by Engle and others are shown to be special cases of a general four-dimensional VAR. Many of the restrictions that the traditional models embody are shown to be largely rejected by the data in favor of models with greater parameterization. This would seem to be due, at least in the short run, due to the fact that the traditional models try to track local sectoral fluctuations by using broader aggregates. This implicitly minimizes the role of local productivity shocks, which, according to the variance decomposition, turn out to be quite important. In the long run there is some connection between local sectoral movements and broader aggregates via cointegrating relationships, but the relationship is not homogenous, and the imposition of shift-share type restrictions is not recommended even in the long run.

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Table 1
Unit root tests

	US	Philadelphia	Dallas	Atlanta	Chicago	Los Angeles
Total	-1.88	-2.69	-3.92*	-2.86	-3.01	-3.74*
Construction	-3.79*	-3.20	-3.37	-3.01	-1.86	-3.48*
Durable Manufacturing	-1.60	-2.81	-1.53	-0.59	-2.75	-3.49*
Nondurable Manufacturing	-1.35	-2.14	-2.51	-0.78	-1.60	-2.06
Trade, Tranportation and Utilities	-1.55	-3.42	-3.34	-2.51	-2.91	-2.91
Finance	-3.45*	-2.52	-2.66	-1.05	-2.97	-2.09
Information	-1.39	-0.23	0.10	0.61	-0.57	-1.65
Professional and Business Services	-2.10	-3.19	-3.16	-2.17	-2.66	-2.16
Education and Health Services	-2.14	-2.07	-2.39	-1.46	-3.22	-3.34
Leisure and Hospitality Services	-1.00	-2.40	-1.55	-1.58	-1.97	-2.24
Other Services	-0.79	-3.22	-0.55	-1.82	-2.61	-2.27
Government	-1.69	-1.81	-2.47	-2.97	-.059	-3.04

The table entries are the t-values from an Augmented Dickey-Fuller test for unit roots in the indicated series. Asterisks indicate a test-statistic with a prob-value between one and five percent for rejecting the null hypothesis that a unit root exists, against the stationary alternative. The Dickey-Fuller regressions contain an intercept, a time trend, and lags of the first difference as selected by the Schwarz information criterion.

Table 2
Trace Tests of the Long-Run Shift Share

	Philadelphia	Dallas	Atlanta	Chicago	Los Angeles
Construction	3	3	3	3	2
Durable Manufacturing	1	2	3	3	2
Nondurable Manufacturing	3	2	1	2	2
Trade, Transportation and Utilities	2	1	0	1	3
Finance, Insurance and Real Estate	1	1	2	1	2
Information Services	1	1	1	1	3
Professional and Business Services	2	2	2	1	2
Education and Health Services	3	2	2	3	3
Leisure and Hospitality Services	2	2	2	1	3
Other Services	1	1	2	1	3
Government	3	3	2	1	3

The table entries are the number of cointegrating vectors in a 4 equation system consisting of the logs of national employment, total city employment, total industry employment and city-industry employment for the indicated row and column. Sequential trace tests were employed to reject (or fail to reject) ranks in the A_0 matrix of zero, one, and two at 5 % critical values.

Table 3
Trace Tests of the Constant Share Model

	Philadelphia	Dallas	Atlanta	Chicago	Los Angeles
Total	24.39* ≠1	14.73	23.1*9 ≠1	13.76	42.12* =1
Construction	57.24* ≠1	18.40* ≠1	13.58	87.28* ≠1	25.20* ≠1
Durable Manufacturing	7.27	8.51	9.98	15.53* ≠1	30.88* ≠1
Nondurable Manufacturing	13.29	5.87	5.01	22.69* ≠1	7.79
Trade, Transportation and Utilities	29.95 ≠1	21-28 ≠1	22.10 ≠1	12.31	13.45
Finance, Insurance and Real Estate	10.61	6.39	11.34	18.39 *≠1	23.07* ≠1
Information Services	2.87	11.98	12.81	7.85	12.90
Professional and Business Services	11.22	39.92* ≠1	15.74* ≠1	26.85* ≠1	14.23
Education and Health Services	35.34* ≠1	19.69* =1	19.78* =1	20.87* =1	28.66* =1
Leisure and Hospitality Services	35.43 *=1	62.67* ≠1	61.21 * ≠1	120.41* ≠1	13.43
Other Services	19.31* =1	32.32* ≠1	12.81	33.56* ≠1	24.31 ≠1
Government	27.15* ≠1	25.51* ≠1	20.25* ≠1	12.00	8.57

With the exception of the first row, the table entries are the trace test statistic for cointegration between the indicated city industry and its national counterpart. An asterisk indicates significance (i.e. cointegration) at the 5% level. For each significant result, the notation =1 indicates that the normalized cointegration coefficient contains 1 in its 95% confidence intervale, ≠1 indicating the contrary. The first row is the corresponding statistic for aggregate employment.

Table 4
Trace tests of the Multiplier Model

	Philadelphia	Dallas	Atlanta	Chicago	Los Angeles
Construction	30.54 ≠1	66.18 ≠1	31.19 ≠1	43.39 ≠1	5.44
Durable Manufacturing	5.06	7.62	7.75	9.87	14.49
Nondurable Manufacturing	4.29	36.94 ≠1	4.17	20.69 ≠1	9.80
Trade, Transportation and Utilities	11.84	7.64	8.23	33.90 ≠1	39.20 ≠1
Finance, Insurance and Real Estate	6.24	14.66	10.09	12.77	45.5 ≠1
Information Services	5.83	2.91	6.85	2.76	41.02 ≠1
Professional and Business Services	6.09	41.45 ≠1	30.44 ≠1	25.37 ≠1	48.02 ≠1
Education and Health Services	11.66	3.66	8.42	15.08	10.16
Leisure and Hospitality Services	4.39	11.92	9.16	14.12	19.13 ≠1
Other Services	2.73	10.29	25.3 =1	3.51	8.56
Government	31.83 =1	13.15	7.55	58.80 ≠1	14.92

The table entries are the trace test statistic for cointegration between the indicated city industry and its regional aggregate. An asterisk indicates significance (i.e. cointegration) at the 5% level. For each significant result, the notation =1 indicates that the normalized cointegration coefficient contains 1 in its 95% confidence intervale, ≠1 indicating the contrary.

Table 5
Philadelphia VARs

	Model A				Model B				Model C				Model D			
	n	i	m	e	n	i	m	e	n	i	m	e	n	i	m	e
C	17.6	31.5	17.7	33.2	11.7	35.9	9.2	63.3	11.0	8.1	27.6	53.4	10.6	9.9	20.3	59.3
DM	39.8	21.0	20.1	19.0	13.0	9.5	26.6	50.9	21.2	61.3	9.8	7.8	60.3	20.9	1.7	17.1
NM	28.9	22.9	21.8	26.4	3.5	3.6	35.5	57.4	1.5	13.5	28.1	56.9	1.4	9.4	20.5	68.8
TU	24.2	23.9	26.7	25.2	13.3	29.1	10.8	75	2.9	23.9	72.4	13.2	30.8	8.8	48.2	12.2
F	28.8	25.5	22.0	23.7	3.8	26.6	8.0	61.6	8.8	22.0	20.3	48.8	15.2	9.2	22.1	53.5
IS	30.7	25.7	21.4	22.2	4.6	1.9	57.3	36.2	17.3	50.8	2.2	29.7	31.8	37.2	4.2	26.8
PS	26.4	24.3	26.4	22.9	1.4	0.7	58.0	39.9	15.2	33.1	6.4	45.4	24.9	15.0	16.3	43.8
ES	26.5	18.9	36.8	17.8	1.6	0.7	56.1	41.6	4.8	2.3	36.4	56.6	15.3	0.5	36.0	48.2
LS	22.1	23.5	25.0	29.4	0.6	2.2	47.4	49.8	5.3	7.5	39.7	47.5	4.8	10.4	20.4	64.4
OS	24.6	26.5	24.1	24.8	2.4	1.1	31.7	64.8	8.1	5.1	8.1	78.6	1.5	1.5	7.1	89.9
G	28.7	23.7	26.9	20.7	0.7	2.9	23.9	72.4	4.3	3.7	32.2	59.7	10.1	3.4	29.7	56.7

The table entries are the percentage of the 24-month forecast error variance of local employment that can be ascribed to the indicated shock.

Table 6
Dallas VARs

	Model A				Model B				Model C				Model D			
	n	i	m	e	n	i	m	e	n	i	m	e	n	i	m	e
C	53.0	20.6	5.7	20.7	8.5	1.5	31.7	62.4	63.1	1.2	20.2	15.5	42.7	9.9	21.9	25.5
DM	72.3	19.6	3.2	4.9	22.7	6.8	10.8	59.7	56.3	33.9	0.7	9.1	58.0	25.8	2.7	13.6
NM	54.4	31.1	1.5	13.0	8.0	4.3	52.5	35.3	73.6	12.4	9.6	4.3	75.4	8.1	9.8	6.7
TU	60.5	7.6	15.5	16.4	3.7	2.4	54.9	39.0	49.5	3.4	44.5	2.6	55.7	4.9	28.2	11.2
F	69.5	24.1	2.3	4.1	6.7	6.8	22.0	64.5	32.1	19.5	2.5	45.9	26.2	17.4	23.1	33.2
IS	63.8	18.5	7.7	10.0	21.3	4.5	4.8	69.4	59.0	6.9	1.3	32.8	58.0	2.8	1.4	37.8
PS	9.7	68.9	9.8	11.6	5.3	4.0	59.8	30.9	22.8	59.3	9.7	8.2	40.5	28.7	17.6	13.2
ES	12.2	48.3	7.3	32.2	1.2	1.6	55.4	41.9	1.2	0.8	10.4	87.6	1.3	13.8	32.3	52.5
LS	64.5	14.9	6.7	13.9	1.0	1.3	36.2	61.4	45.5	7.5	18.4	28.6	34.5	12.0	16.9	36.6
OS	66.1	18.9	5.0	10.0	3.5	1.2	20.0	75.3	38.4	4.4	7.6	49.6	37.7	8.0	6.7	47.5
G	52.9	22.3	10.3	14.5	0.7	1.7	14.3	83.4	19.6	8.2	24.2	48.1	12.5	7.4	22.4	57.7

The table entries are the percentage of the 24-month forecast error variance of local employment that can be ascribed to the indicated shock.

Table 7
Atlanta VARs

	Model A				Model B				Model C				Model D			
	n	i	m	e	n	i	m	e	n	i	m	e	n	i	m	e
C	27.7	25.9	22.7	23.8	6.7	3.7	39.2	50.4	59.1	5.1	7.3	28.6	27.0	14.3	29.2	29.5
DM	27.1	25.5	22.5	24.8	5.1	4.0	14.5	76.4	30.8	24.8	13.1	31.2	44.4	5.4	3.2	47.1
NM	25.4	27.2	22.6	24.8	2.1	3.0	22.5	72.3	15.5	1.6	33.0	49.9	49.9	9.4	9.6	31.2
TU	19.1	30.3	17.8	32.8	4.3	2.4	55.7	37.5	4.3	2.4	55.7	37.5	57.0	2.7	20.4	19.9
F	27.9	24.8	23.9	23.4	2.5	5.7	29.2	62.6	39.3	6.3	3.5	50.9	33.5	4.8	10.3	51.4
IS	35.7	25.2	17.7	21.4	8.7	10.7	14.2	66.4	31.1	38.4	22.2	8.3	40.1	34.5	5.6	19.8
PS	25.9	24.8	27.5	21.8	3.9	4.6	65.7	25.8	50.0	18.7	14.6	16.7	46.8	16.5	4.5	32.1
ES	22.8	22.5	28.4	26.3	1.9	0.9	35.8	61.4	6.5	5.3	14.9	73.3	0.9	19.5	28.8	50.8
LS	22.0	25.6	25.2	27.2	1.6	1.2	45.0	52.2	22.0	17.6	22.8	37.7	27.1	17.4	17.5	38.0
OS	26.8	23.9	25.0	24.3	3.7	0.3	6.1	89.9	28.4	1.1	2.6	67.9	16.9	3.2	13.3	66.6
G	30.9	19.2	29.0	27.9	23.7	27.7	20.7	77.3	10.2	18.6	19.2	51.9	1.4	14.0	30.2	54.4

The table entries are the percentage of the 24-month forecast error variance of local employment that can be ascribed to the indicated shock.

Table 8
Chicago VARs

	Model A				Model B				Model C				Model D			
C	45.1	8.8	38.9	16.8	30.0	14.8	38.4	62.3	15.6	2.4	33.0	49.0	16.2	4.6	31.0	48.2
DM	36.4	24.9	19.6	19.1	9.7	11.3	32.8	46.3	25.0	17.4	2.2	55.4	17.4	2.2	55.4	57.2
NM	35.0	23.4	22.1	19.6	4.9	4.8	47.8	42.6	21.1	63.3	4.8	10.8	15.7	48.9	8.2	27.2
TU	17.7	30.0	18.4	34.0	2.6	2.0	51.7	43.7	47.6	1.4	27.8	23.1	46.4	1.3	32.2	20.1
F	28.0	24.3	22.8	25.0	2.0	1.9	30.1	66.1	3.4	2.2	40.4	54.0	7.5	1.3	6.7	84.6
IS	28.4	23.3	28.6	33.5	22.9	21.8	21.8	77.1	24.7	30.2	2.2	42.8	10.1	23.9	9.5	56.5
PS	21.3	29.6	18.3	30.9	3.4	2.4	56.7	37.6	18.7	45.5	16.9	18.9	51.4	17.0	13.9	17.7
ES	36.3	13.6	33.9	16.1	1.0	3.6	45.7	49.6	10.0	42.8	14.4	32.8	4.1	17.4	20.7	57.8
LS	13.5	31.5	14.3	40.7	1.0	1.2	43.1	54.7	23.8	10.2	34.4	31.6	21.0	10.8	33.9	34.3
OS	25.0	25.3	24.0	25.7	2.0	1.5	29.9	66.6	19.6	15.1	27.9	37.4	35.0	4.5	22.1	38.3
G	30.2	19.4	31.7	18.6	0.6	1.1	29.9	68.5	2.7	5.2	47.2	44.9	6.0	6.2	41.1	46.6

The table entries are the percentage of the 24-month forecast error variance of local employment that can be ascribed to the indicated shock.

Table 9
Los Angeles VARs

	Model A				Model B				Model C				Model D			
	n	i	m	e	n	i	m	e	n	i	m	e	n	i	m	e
C	32.0	21.9	26.6	19.5	3.5	2.8	48.2	45.5	20.7	17.7	5.9	55.6	26.1	6.7	4.9	62.4
DM	36.6	23.1	22.3	18.0	8.1	6.6	45.6	39.7	17.6	43.3	5.9	33.3	29.3	46.8	3.5	20.5
NM	5.5	5.4	51.2	33.6	23.2	27.9	15.4	37.9	31.0	12.5	1.5	55.0	23.5	24.2	3.9	48.4
TU	22.2	27.3	22.5	28.0	0.8	2.2	66.9	30.0	30.9	0.2	58.5	10.4	24.1	2.2	55.4	18.3
F	25.1	25.3	26.3	23.4	2.4	3.1	40.2	54.3	12.9	5.5	13.6	68.0	6.0	20.4	14.4	59.2
IS	26.9	19.2	33.1	20.8	4.7	2.2	23.1	70.0	22.7	22.6	7.5	47.3	6.0	7.3	4.4	82.3
PS	27.4	24.2	27.3	21.1	3.5	2.4	54.4	39.7	35.0	6.0	8.3	50.6	33.0	6.0	10.0	51.1
ES	21.8	30.5	21.2	26.4	0.1	0.6	30.0	69.3	0.8	16.1	8.5	74.6	4.6	9.9	8.5	77.1
LS	25.2	24.5	24.9	25.3	0.3	1.9	53.1	44.7	9.8	0.4	9.2	80.5	16.7	19.4	22.7	41.2
OS	27.1	23.6	29.4	20.0	0.3	2.3	58.0	39.3	3.9	0.4	7.9	87.8	4.0	2.3	26.4	67.3
G	25.8	25.5	24.2	24.5	2.0	2.1	19.6	76.3	0.4	1.1	17.9	80.6	6.5	5.1	15.0	73.3

The table entries are the percentage of the 24-month forecast error variance of local employment that can be ascribed to the indicated shock.