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# The Long-Run Relationship between House Prices and Income: Evidence from Local Housing Markets

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Many in the housing literature argue that house prices and income are cointegrated. I show that the data do not support this view. Standard tests using 27 years of national-level data do not find evidence of cointegration. However, standard tests for cointegration have low power, especially in small samples. I use panel-data tests for cointegration that are more powerful than their time-series counterparts to test for cointegration in a panel of 95 metro areas over 23 years. Using a bootstrap approach to allow for cross-correlations in city-level house-price shocks, I show that even these more powerful tests do not reject the hypothesis of no cointegration. Thus the error-correction specification for house prices and income commonly found in the literature may be inappropriate.

In the second half of the 1980s, real house prices rose about 3% per year.[1] Then in 1990 alone, real prices tumbled almost 5%. Prices continued to fall, on balance, through the end of 1994, reversing more than half of the gains posted in the late 1980s; prices did not return to their 1989 level until almost 10 years later. Many coastal cities experienced even wilder swings: Real house prices rose about 65% in Los Angeles and about 45% in Boston in the second half of the 1980s, only to fall 30% in Los Angeles and 20% in Boston during the next 5 years.

Many real-estate market observers think that housing prices got too far ahead of fundamentals in the 1980s—especially in many coastal cities—and that the poor performance of house prices during the first half of the 1990s was the inevitable aftermath. Real house prices have risen rapidly in recent years, sparking fears that the housing market is once again overvalued.

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[1] All house price measures in this article are from the Office of Federal Housing Enterprise Oversight's (OFHEO) weighted repeat-sales price index unless otherwise noted. Nominal prices are deflated by the personal consumption deflator from the National Income and Product Accounts.

**Figure 1** ■ Ratio of house prices to per capita personal income.


*Note:* The index is the ratio of the Office of Housing Enterprise Oversight repeat-sales price index divided by per capita personal income as reported by the Bureau of Economic Activity.

Of particular concern to many is the fact that house-price gains have dwarfed per capita income gains since 1998. As can be seen in Figure 1, the ratio of house prices to per capita personal income moved up rapidly beginning in the late 1990s after trending down for most of the previous 20 years.[2] More specifically, from the middle of 1997 to the middle of 2002, real house prices rose about 28% while real per capita personal income rose about 15%. In contrast, during the previous 20-year period, real house prices rose only 8% while real per capita income rose 35%. The recent performance of house prices relative to income is taken as evidence by some that house prices are out of line with "fundamentals" and that prices must stagnate or fall to allow income to catch up.

This idea is commonly formalized in the housing literature by positing a cointegrating relationship between house prices and fundamentals such as income and then estimating an error-correction specification (Abraham and Hendershott 1996, Malpezzi 1999, Capozza *et al.* 2002, Meen 2002). That is, house prices and income are thought to be linked by a stable long-run relationship; they

[2] In this article, I focus on per capita income rather than household income for two reasons. First, household size—the link between per capita and household income—is endogenous to the household formation problem and therefore to housing demand. Second, I do not have city-level data on the number of households.

may drift apart temporarily, but their tendency is to return to their long-run equilibrium.

The purpose of this article is to test this view of the housing market. If prices and income are cointegrated, then the gap between the two may be a useful indicator of when house prices are above or below their equilibrium values and therefore a useful predictor of future house-price changes. Conversely, if prices, income and other fundamentals such as interest rates and construction costs are not cointegrated, then the error-correction specifications common in the literature are inappropriate, and house prices need not stagnate or fall just because they have grown more quickly than has income of late.[3]

Many researchers simply assume that house prices and fundamentals are cointegrated (Abraham and Hendershott 1996, Capozza *et al*. 2002). Others implicitly assume that they are not (Poterba 1991). Meen (2002) is the only study I am aware of that tests for cointegration of prices and fundamentals using national-level data. His reported tests do not find evidence for cointegration at conventional significance levels. However, Meen argues that the test statistics are "near" their critical values, and therefore he concludes that prices and fundamentals are cointegrated. One contribution of this article is to show that using 27 years of national-level data, one does not find evidence that prices, income and other fundamentals are cointegrated. Thus my results, if not my interpretation, are in accord with Meen, and they suggest that it is inappropriate to model house-price dynamics using an error-correction specification.

However, cointegration tests are known to have low power, particularly in small samples (Banerjee 1999). A time span of 27 years may be too short to estimate what may be a genuine long-run relationship with slow adjustment. Starting with Quah (1990) and Levin and Lin (1992), researchers have developed panel tests for unit roots and cointegration that are more powerful than their standard time-series counterparts. The second, and main, contribution of this article is to apply the recently developed tests of Pedroni (1999) and Maddala and Wu (1999) to a panel of 95 U.S. cities over 23 years. I show that even these more powerful tests cannot reject the hypothesis that prices and income are not cointegrated. My results contradict those of Malpezzi (1999), which found that one *can* reject the null of no cointegration in a similar panel. However, Malpezzi used a panel unit root test, which overstates the likelihood of cointegration because it ignores the first-stage estimation in his residuals-based cointegration test.

[3] The absence of cointegration does not preclude house-price bubbles. For example, one might not find cointegration, which is a linear relationship, if the housing market is beset by nonlinear rational bubbles.

The rest of this article is organized as follows. In the following section, I briefly describe a simple stock-flow model of housing. The purpose of the model is to motivate why prices, per capita income and perhaps other variables might be cointegrated. I then describe the national-level data and show that there is little evidence for cointegration using Engle and Granger's (1987) augmented Dickey–Fuller (ADF) test. City-level tests form the heart of the article. I describe several tests for cointegration in panel data. Three are from Pedroni (1999): one is a panel version of Phillips and Ouliaris' (1990) variance ratio test, one is a panel version of their $Z_\alpha$ test and one is a panel version of Engle and Granger's (1987) ADF test. My fourth test is a Maddala and Wu (1999) version of Engle and Granger's ADF test. I also describe a bootstrap approach that allows for arbitrary cross-correlations of the city-level shocks. I describe the city-level data and show that, using the bootstrapped critical values, none of the tests rejects the null of no cointegration. A secondary result is that the correlations among local housing markets can have a large effect on the tests.

## House Prices and Fundamentals in a Stock-Flow Model of Housing

House prices and fundamentals like income—or some transformation of them—will be cointegrated if they are linked by a long-run equilibrium relationship. A simple stock-flow model of housing similar to those in Poterba (1984) and Topel and Rosen (1988) can provide a theoretical basis for such a relationship. In the model, the service flow from housing is proportional to the stock of housing, $K_t$, and the demand for housing services can be written as

$$R_t = -\alpha K_t + \theta_t, \tag{1}$$

where $R_t$ is the rental rate for a unit of housing and $\theta_t$ is a vector of demand shifters. For simplicity, I assume that $\theta_t$ follows a random walk:

$$\theta_t = \theta_{t-1} + u_t.$$

According to the simple textbook model of house prices, the user cost of housing should equal rent. Ignoring taxes, maintenence and the risk premium,

$$R_t = P_t - \frac{E_t P_{t+1}(1-\delta)}{1+i}, \tag{2}$$

where $P_t$ is the price of housing at time $t$, $E_t$ is the expectations operator and $\delta$ and $i$ are the time-invariant rates of depreciation and interest.[4] Substituting Equation (2) into (1) yields

$$P_t = -\alpha K_t + \theta_t + \beta E_t P_{t+1}, \tag{3}$$

[4] I assume that the rates of depreciation and interest are constant over time to keep the model simple. One could extend the model to allow for time variation in these variables.

where $\beta = (1 - \delta)/(1 + i)$. I assume that $K_t$ is fixed in the short run and that new investment is governed by

$$I_t = \psi P_t + \epsilon_t, \tag{4}$$

where $\epsilon$ is a vector of supply shifters. The usual capital accumulation identity,

$K_t = (1 - \delta)K_{t-1} + I_{t-1}$, implies that

$$P_t = \frac{-\alpha\psi L}{1 - (1 - \delta)} P_{t-1} + \beta E_t P_{t+1} + \theta_t + \frac{1}{1 - (1 - \delta L)} \epsilon_{t-1}, \tag{5}$$

where $L$ is the lag operator.

One can show that, for reasonable values for $\alpha$ and $\psi$, the solution to Equation (5) has real roots, one greater than unity and one less than unity. If we assume that $\epsilon_t$ is stationary, then one can show that

$$P_t = a\theta_t + b(L)u_t + d(L)\epsilon_t, \tag{6}$$

where $b(1) < 0$ and $d(1) < 0$. Thus, in the model, house prices are cointegrated with demand shifters in $\theta_t$. More generally, if elements of $\epsilon_t$ have unit roots, one can show that, in the model, house prices, demand shifters and supply shifters are cointegrated.

One can easily imagine many fundamentals that one might include in $\theta_t$ and $\epsilon_t$—income, population, wealth, construction wages and interest rates come to mind. However, the model does not show that house prices *must* be cointegrated with these fundamentals. Indeed, this simple model illustrates that there are many reasons why such a cointegrating relationship need not exist. For instance, the price elasticity of new investment may not be stable over time because of changes in regulatory conditions or the price elasticity of demand may not be stable because of changing demographics. The model does show what kind of assumptions are needed to generate a cointegrating relationship. Whether one exists is an empirical question.

## National-Level Tests for Cointegration

In this section, I present tests of cointegration of national-level house prices and various fundamentals. Suppose that the hypothesized cointegrating regression is given by

$$x_{0,t} = \alpha + \delta t + \sum_{m=1}^{M} \beta_m x_{m,t} + e_t, \tag{7}$$

where $m = 1, \ldots, M$ indexes I(1) variables and $t = 1, \ldots, T$ indexes time. If the residual $e_t$ is stationary, then we say that the $x$'s are cointegrated. Here, I use the common two-step procedure for testing for cointegration suggested by Engle and Granger (1987), sometimes called an augmented Engle-Granger (AEG) $\tau$ test. In the first stage, I estimate Equation (7) by ordinary least squares to get $\hat{e}_t$. In the second stage, I conduct an ADF $\tau$ test on the residuals. The critical values differ from those of the standard ADF test because the residuals are estimated in the first stage.

My source for house-price data is the repeat-sales price index for existing homes, which is published by the OFHEO.[5] The index is based on price changes for homes that are resold or refinanced, but it does not control for changes to the house through improvement or neglect; that is, while it does hold some characteristics constant, it is not a true quality-adjusted price. In addition, the repeat-sales sample excludes homes with jumbo, Federal Housing Administration, or Department of Veterans Affairs mortgages (Calhoun 1996). I used the Bureau of Economic Activity's (BEA) measure of total personal income, the Census Bureau's measure of population and the Bureau of Labor Statistics' measure of average hourly wages for construction workers. I used the Standard and Poor's 500 stock index to measure stock-market wealth. I included the level of the personal consumption deflator from the BEA to control for inflation. All data are quarterly.

To calculate the user cost of housing, I used a weighted average of the rates on fixed-rate and 1-year adjustable-rate contracts for 30-year loans; the weights were the origination shares.[6] I set expected capital gains to the average percentage increase in the house-price index during the previous 3 years.

The standard unit root tests do not reject the hypotheses that house prices, per capita income, population, the stock market and construction wages all have a unit root, but that the user cost does not.[7]

Table 1 displays AEG $\tau$ tests for cointegration of several sets of variables using quarterly data from 1975:Q1 to 2002:Q2. The first-stage levels equation

[5] Alternative tests based on the average existing house price from the National Association of Realtors yielded similar results.

[6] My measures of federal and state and local tax rates are from the FRB/US model. See Reifschneider, Tetlow and Williams (1999) for more information about the FRB/US model.

[7] The ADF test statistics based on 109 observations are as follows: house prices ($-2.3$), per capita income ($-2.4$), population ($-3.2$), the stock market ($-1.7$), construction wages (0.3) and the user cost ($-3.9$). All tests include a time trend and five autoregressive terms. The 5% critical value is about $-3.4$.

**Table 1** ■ National-level tests for cointegration of house prices and fundamentals.

| Independent variables (log values) | First-Stage Levels Regressions dependent variable is log(price) 1 | 2 | 3 | 4 |
|---|---|---|---|---|
| | Null of no cointegration, 1975:Q1 to 2002:Q1 | | | |
| Per capita income | 0.70** | 1.45** | 1.57** | 1.08** |
| | (0.05) | (0.20) | (0.15) | (0.16) |
| Population | — | −2.94** | — | −15.8** |
| | | (0.76) | | (2.47) |
| User cost | — | — | — | −0.03** |
| | | | | (0.004) |
| Stock market | — | — | — | −0.06** |
| | | | | (0.02) |
| Construction wage | — | — | — | 1.43** |
| | | | | (0.19) |
| PCE deflator | 0.03 | −0.38** | −0.76** | 0.250 |
| | (0.10) | (0.14) | (0.15) | (0.20) |
| Trend × 10 | — | — | 0.74** | 0.02** |
| | | | (0.12) | (0.01) |
| | Second-Stage Test Results | | | |
| AEG $\tau$ stat | −2.2 | −2.0 | −1.9 | −3.1 |
| Critical value (10%) | −3.5 | −3.8 | −3.8 | −5.0 |

**Significant at 0.05. Standard errors are in parentheses. AEG = Augmented Engle-Granger. PCE = Personal Consumption Expenditures.

yields an estimate of the cointegrating vectors, which are shown in the upper panel of the table. The coefficient estimates indicate that per capita income and the construction wage have positive effects on house prices and that the user cost has a negative effect. Surprisingly, population and the stock market have negative effects. Of course, if there is no cointegrating relationship, these levels regressions are spurious. The lower panel presents the second-stage tests with 10% critical values that are based on the dimension of the proposed cointegrating vector and the presence or absence of a time trend (Davidson and MacKinnon 1993). The lower panel of Table 1 shows that in none of the cases can we find strong evidence for the cointegration of house prices and fundamentals. In other words, the national-level data do not support the view that the log levels of house prices and various fundamentals are linked by a long-run stationary relationship.[8]

One criticism of these tests is that they are known to have low power against the alternative of cointegration, particularly when, as is the case here, the sample

[8] Meen (2002) conducted similar tests and concluded that prices and fundamentals are cointegrated. However, his reported test statistics were quite far from conventional critical values.

size is small. Thus the evidence in Table 1 may not be convincing to those who have strong priors that house prices and fundamentals, particularly income, are cointegrated.

## City-Level Tests for Cointegration

Quah (1990, 1994) and Levin and Lin (1992) were among the first to devise panel tests for unit roots and to show that they can offer a substantial improvement in power relative to separate tests for each cross-sectional unit of the panel.[9] The literature has blossomed since then. Some, like Im, Pesaran and Shin (1997), have devised tests that impose fewer restrictions than did Levin and Lin. Others, like Pedroni (1997, 1999, 2001), developed related tests for cointegration. Banerjee (1999) provides an overview of the literature.

In this section, I briefly describe three panel cointegration tests of Pedroni (1999) and of Maddala and Wu (1999). One of the underlying assumptions of all the tests is that shocks are either independent across cross sections or that the cross-sectional dependence can be modeled as an aggregate time effect. For the purposes of this article, that assumption implies that shocks to housing markets in, say, San Francisco and Seattle have the same correlation as shocks to housing markets in Philadelphia and New York. As this is an unattractive feature, I discuss bootstrapped versions of the tests that relax this assumption by allowing for the cross-sectional dependence among cities evident in the data. I then describe the city-level data and present the test results.

### *Some Panel Cointegration Tests*

Suppose that the hypothesized cointegrating regression for each city is given by

$$x_{0,i,t} = \phi_i + \delta_i t + \sum_{m=1}^{M} \beta_{m,i} x_{m,i,t} + e_{it}, \tag{8}$$

where $i = 1, \ldots, N$ indexes the city, $m = 1, \ldots, M$ indexes variables and $t = 1, \ldots, T$ indexes time. Notice that this specification admits city-specific intercepts and time trends.

In this article, I use Pedroni's panel-data versions of Phillips and Ouliaris' (1990) variance ratio test, $P - v$, Phillips and Ouliaris' $Z_\alpha$ test, $P - Z_\alpha$, and Engle and Granger's (1987) ADF test, $P - AEG$. These residual-based panel tests have the same structure as do their time-series counterparts, but they are constructed

[9] See also Levin, Lin and Chu (2002).

by pooling information from the cross-sectional units.[10] See Appendix A for details on how the pooling is done.

Pedroni (1999) shows that the appropriately standardized version of his test statistics are asymptotically normal.[11] As with the standard time-series cases, under the alternative, the $P - v$ test diverges to positive infinity, so the right tail of the normal distribution is used for rejection, and the $P - Z_\alpha$ and $P - AEG$ tests diverge to negative infinity, so the left tail is used to reject.

Maddala and Wu (1999) present an alternative, and very general, test for cointegration based on Fisher (1932). Maddala and Wu's test is based on averaging the $p$ values for *any* test from each cross-sectional unit.

Suppose that we implement any cointegration test for each cross-sectional unit. Under the null, the significance levels for each test, $p_i$ are distributed uniformly over (0, 1). This implies that

$$-2 \log p_i \sim \chi^2(2).$$

Under the assumption that the tests are independent, the Maddala–Wu (MW) test statistic is

$$MW = -2 \sum_{i=1}^{N} \log p_i \sim \chi^2(2N),$$

and the right tail of the distribution is used for rejection. This test can be applied using any underlying test. However, one must simulate an approximation to the entire distribution of whatever test statistic one is using for each cross-sectional unit in order to calculate the $p_i$'s. The MW tests I present in this article are based on AEG $\tau$ tests, where I simulated the AEG $\tau$ test distributions for a 23-year time series. That is, I did not use the asymptotic distribution of the $\tau$ statistic.

The validity of Pedroni's tests depends in part on the assumption that any cross-sectional correlations are adequately captured by an aggregate time effect; the MW test requires complete independence. The assumption of cross-sectional independence is likey violated in the data because local housing market shocks

[10] Pedroni also presents a panel version of Phillips and Ouliaris' $Z_t$ test. The results from the $P - Z_\alpha$ test and $P - Z_t$ were almost identical, so I only report the $P - Z_\alpha$ test results in this article. In addition, the tests I use here require that the AR(1) terms in the second-stage residual regressions for each city be equal to each other (and less than one) under the alternative hypotheses. Pedroni also proposes versions of his tests that do not impose this equality restriction. The results are not sensitive to the choice of tests.

[11] Pedroni calculated each test statistics' mean and variance by simulation. Pedroni generously gave me the RATS code to construct the test statistics.

will likely be correlated in ways that are not captured by a simple time effect (Calem and Case 2005). One way to address this problem is to bootstrap empirical distributions of the test statistics under the null to calculate critical values for the test. The key is to maintain the cross-sectional dependence while resampling.[12]

More specifically, following Maddala and Wu (1999), one can get the bootstrap sample by estimating

$$\Delta x_{m,i,t} = \eta_{m,i}(L)\Delta x_{m,i,t-1} + u_{m,i,t} \quad t = 1, \ldots, T \tag{9}$$

for each series $m$ in city $i$ and then calculating the residuals, $\hat{u}_{m,i,t}$. Then, let $\psi(t) \sim$ uniform $(1, T)$ index the random resampling and construct

$$\tilde{u}_{m,i,t} = \hat{u}_{m,i,\psi(t)} \quad \forall m, i, t.$$

Note that the same $\psi(t)$ is used for each city $i$, thereby maintaining the cross-correlation that exists in the data.

Given each set of resampled $u$'s, construct

$$\tilde{v}_{m,i,0} = \sum_{j=0}^{J} \hat{\eta}_{m,i}(L)\check{u}_{m,i,-j}$$

$$\tilde{v}_{m,i,t} = \hat{\eta}_{m,i}(L)\tilde{v}_{m,i,t-1} + \tilde{u}_{m,i,t}$$

$$\tilde{x}_{m,i,0} = 0$$

$$\tilde{x}_{m,i,t} = \tilde{x}_{m,i,t-1} + \tilde{v}_{m,i,t},$$

where the $\check{u}_{m,i,-j}$ are drawn as a separate bootstrap sample of size $J$.[13]

The null of no cointegration is true by construction for the resampled data, so one can build up the empirical distribution of the test statistic under the null by replicating the process of resampling and re-estimation. The bootstrapped test results presented later in this article are based on 20,000 replications of 23 "years" of data for each city, where $J = 100$. These bootstrapped distributions are, by construction, the small-sample distributions and are therefore not directly comparable to the asymptotically normal distributions of Pedroni's tests, but are comparable to the MW test. To help with comparisons, I also

[12] One could also use a spatial-temporal autoregressive model that incorporates cointegration tests (Pace, Barry and Sirmans 1998, Pace, Barry, Gilley and Sirmans 2000). However, this is beyond the scope of this article.

[13] This is equivalent to setting $\tilde{u}_{m,i,0} = 0$ and simulating $J + T$ observations but only using the last $T$ observations.

constructed the small sample distributions for Pedroni's test statistics under the assumption of cross-sectional independence. That is, for each city I simulated 20,000 replications of 23 "years" of data for three unit-root processes under the assumptions of no cointegration and no cross-city correlations, and then I tabulated the values of Pedroni's test statistics.

### *The City-Level Data*

My data sources at the city and national levels are the same. OFHEO publishes a quarterly repeat-sales price index for over 300 metropolitan areas. However, only a short-time sample exists for many of these cities. For the purposes of this article, I restricted my attention to a sample of 95 cities for which the price data began in 1978. See Appendix B for a list of the cities. Total personal income at the city level is available only at an annual frequency.[14]

Figures 2 through 6 display house prices and per capita income for 15 cities, all in current dollars.[15] I chose to display these cities because they provide a good picture of the behavior of prices and income in most U.S. cities. One can see that, although both prices and per capita income have trended up over time, their levels can diverge for many years. In addition, the charts show that coastal cities have seen more pronounced swings in prices. It is not at all clear from these pictures that house prices and per capita income are cointegrated.

### *Results and Interpretation*

The city-level cointegration tests are based on first-stage cointegrating relationships of house prices on per capita income and population. I estimated the first-stage regressions, which all include a time trend, separately for each city. It is important to note that the bootstrap approach can account for aggregate effects due to factors, such as monetary or fiscal policy, national economic conditions, changes to the housing-finance industry or cost shocks. Thus, I am really testing for the cointegration of prices, income and population conditional on these other, potentially nonstationary, factors.

Table 2 displays the main results of the article. The first column contains the value of each test statistic, the second column contains their asymptotic

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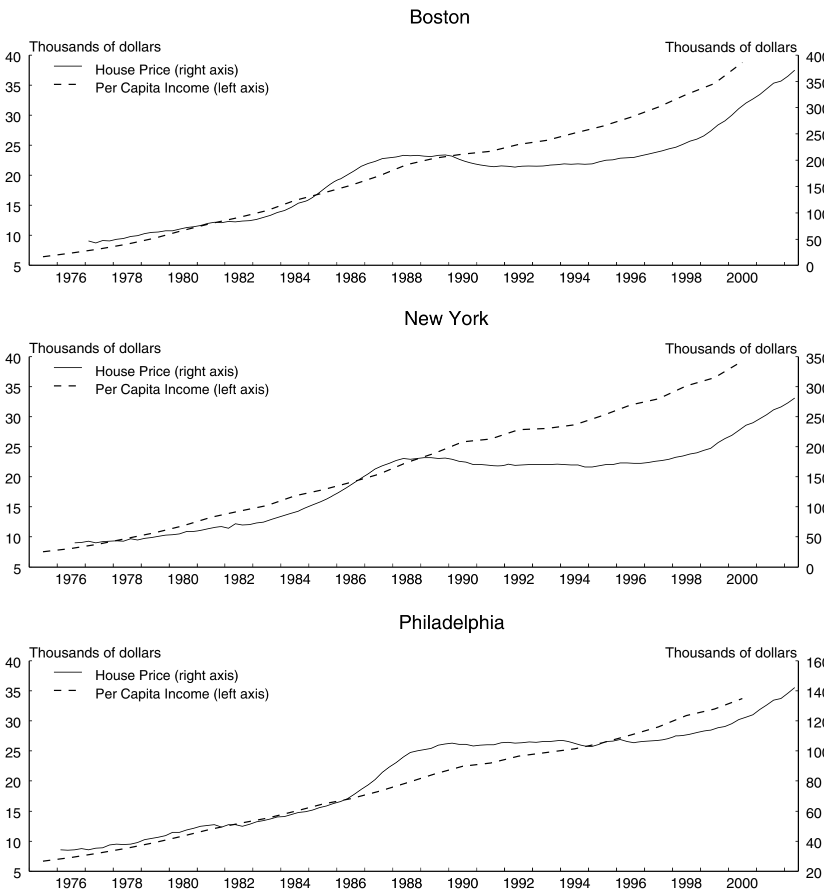
[14] Shiller and Perron (1985) show that the power of unit root tests depends more on the number of years covered by the data set than on the number of observations. Thus, the cost in terms of power of using annual data for the city-level analysis is likely small.

[15] I calculated the current-dollar values by multiplying the median house price in 2001:Q4 published by the National Association of Realtors by the OFHEO index, reindexed to that quarter, for each city.

**Figure 2** ■ House prices and per capita income.

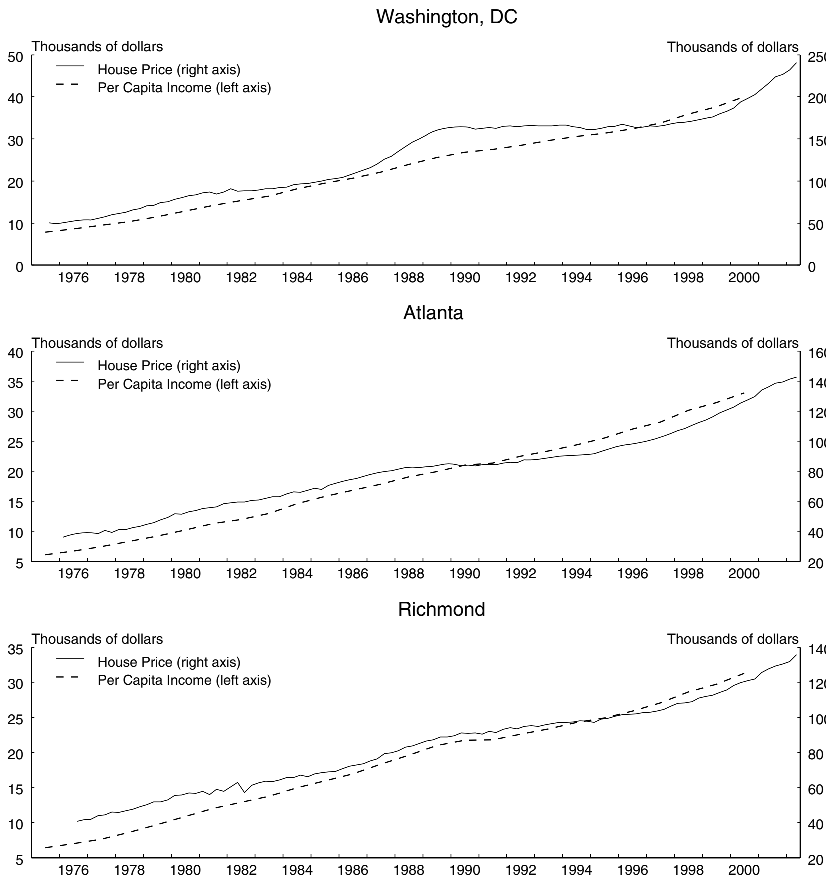



*Note:* House prices are the Office of Federal Housing Enterprise Oversight repeat-sales index multiplied by the nominal median sales price in 2001:Q4 from the National Association of Realtors. Per capita income at the city level is reported by the Bureau of Economic Activity as an annual average through 2000.

$p$ values, the third column contains their small-sample $p$ values and the fourth column contains their bootstrapped $p$ values. The conclusions one might draw from the asymptotic $p$ values are in severe conflict. The $P - v$ and $P - AEG$ tests alone (lines 1 and 3) seem to provide strong evidence for cointegration, while the $P - Z_\alpha$ test (line 2) seems to provide none.

Using the small-sample $p$ values, the case for cointegration looks much weaker. One can no longer reject the null using the $P - AEG$ test, and one cannot reject using the MW statistic. However, one can still reject using the $P - v$ statistic.

**Figure 3** ■ House prices and per capita income.

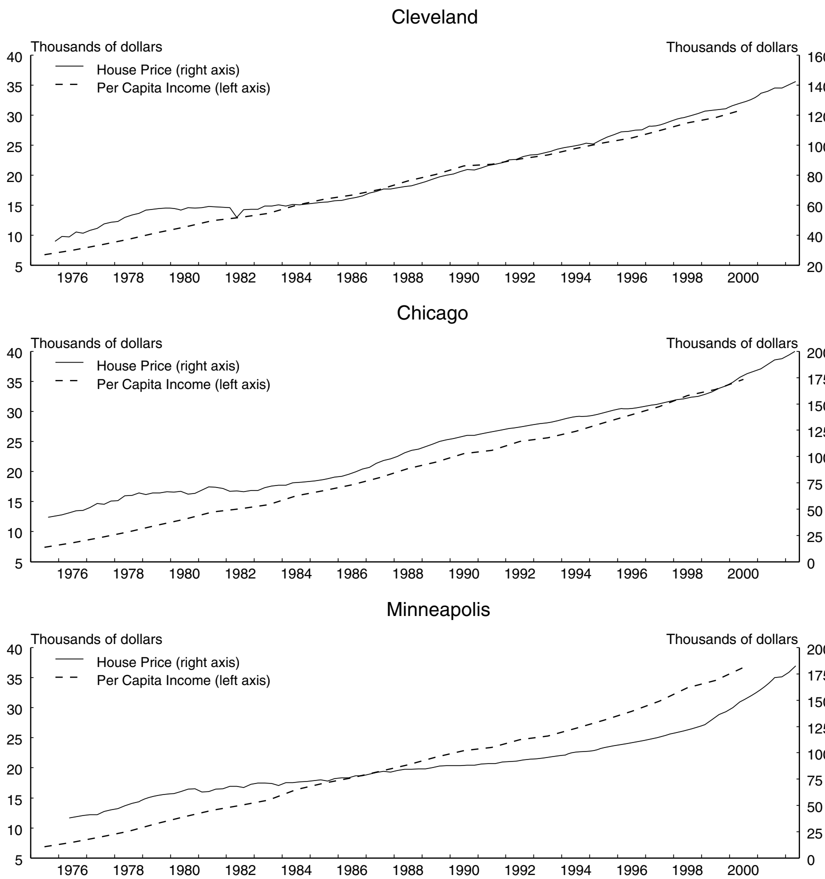



*Note:* House prices are the Office of Federal Housing Enterprise Oversight repeat-sales index multiplied by the nominal median sales price in 2001:Q4 from the National Association of Realtors. Per capita income at the city level is reported by the Bureau of Economic Activity as an annual average through 2000.

Notice that the $P - AEG$ test clearly suffers from extreme size distortion in small samples. Pedroni (2001) finds a less extreme but still significant distortion in the case where $N = 20$ and $T = 40$. My results indicate that the distortions become much worse as $T$ gets even smaller.

Column 4 shows the bootstrapped $p$ values. Note that allowing for arbitrary cross-correlations has a significant effect on the simulated distribution of the test statistics. The $p$ values for the $P - Z_\alpha$, $P - AEG$ and $MW$ statistics are all smaller when one uses the bootstrapped distributions instead of the

**Figure 4** ■ House prices and per capita income.

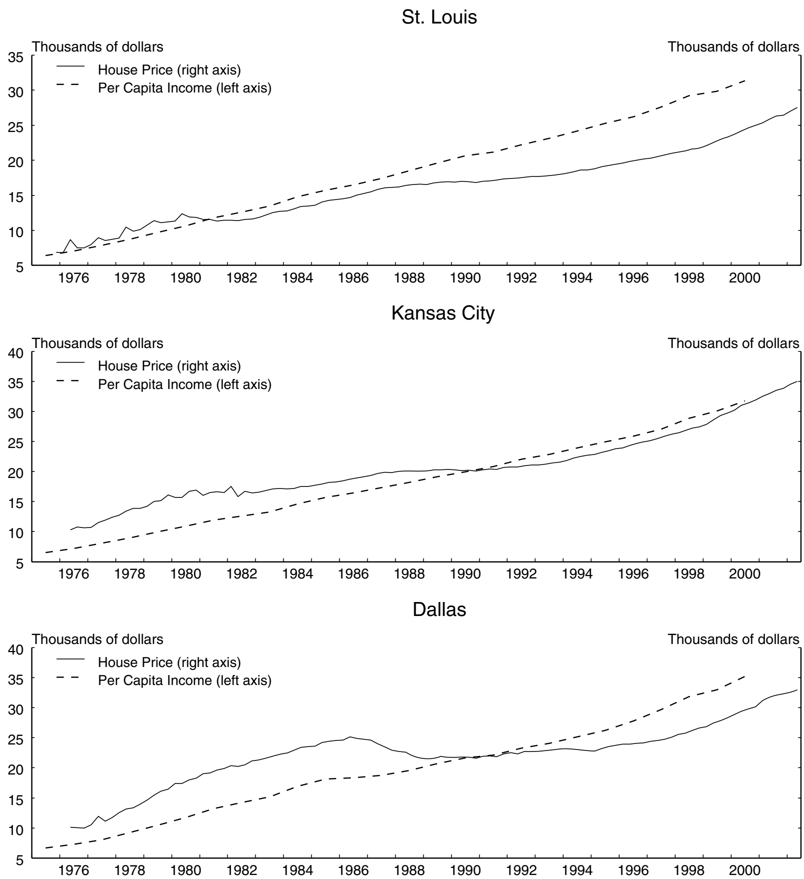



*Note:* House prices are the Office of Federal Housing Enterprise Oversight repeat-sales index multiplied by the nominal median sales price in 2001:Q4 from the National Association of Realtors. Per capita income at the city level is reported by the Bureau of Economic Activity as an annual average through 2000.

small-sample distributions, but one would still not reject the null of no cointegration using conventional significance levels. In contrast, the $p$ value for the $P - v$ statistic is larger using the bootstrapped distribution, and the difference is enough to change the results of the test from rejection to no rejection of the null.

Thus, using conventional significance levels, the tests are all in accordance: House prices, income and population do not appear to be cointegrated at the national or local levels. These results contradict those of Malpezzi (1999),

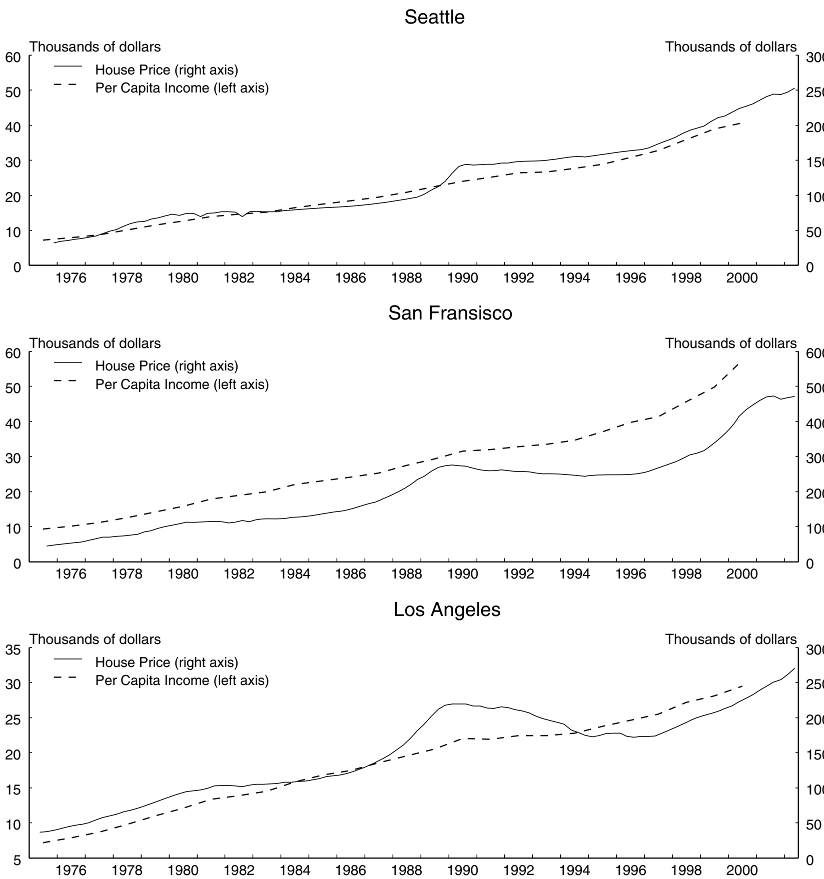
**Figure 5** ■ House prices and per capita income.




*Note:* House prices are the Office of Federal Housing Enterprise Oversight repeat-sales index multiplied by the nominal median sales price in 2001:Q4 from the National Association of Realtors. Per capita income at the city level is reported by the Bureau of Economic Activity as an annual average through 2000.

who found that one *can* reject the null of no cointegration in a similar panel. Malpezzi was able to reject the null of a unit root in the residuals of city-level cointegrating regressions of prices and income using a Levin–Lin test. His interpretation was that prices and income are therefore cointegrated. However, the Levin–Lin test is inappropriate in this context because the critical values tabulated by Levin and Lin (1992) are not adjusted for the fact that Malpezzi's residuals are estimated. In other words, the Levin–Lin test is a unit root test, not a residuals-based cointegration test. While the two types of tests are clearly related, the Levin–Lin test is only applicable to the case of cointegration when

**Figure 6** ■ House prices and per capita income.




*Note:* House prices are the Office of Federal Housing Enterprise Oversight repeat-sales index multiplied by the nominal median sales price in 2001:Q4 from the National Association of Realtors. Per capita income at the city level is reported by the Bureau of Economic Activity as an annual average through 2000.

the researcher imposes the cointegrating relationship. In addition, Malpezzi did not correct for the fact that the Levin–Lin test requires independence across local housing markets.

If house prices and fundamentals are not cointegrated, how should we interpret the error-correction type models of Abraham and Hendershott (1996), Malpezzi (1999) and Capozza *et al*. (2002)? All three find that house prices increase more slowly when actual house prices are above a measure of the long-run equilibrium price level, and all three base their equilibrium measure on a first-stage levels

**Table 2** ■ City-level test for cointegration of house prices, per capita income and population.

| | Test Statistic | Asymptotic $p$ Value | Small-Sample $p$ Value | Bootstrapped $p$ Value |
|---|---|---|---|---|
| | Null of no cointegration; $N = 95$; $T = 23$ | | | |
| 1. $P - v$ | 2.7 | 0.00 | 0.01 | 0.15 |
| 2. $P - Z_\alpha$ | 4.1 | 1.0 | 1.0 | 0.31 |
| 3. $P - AEG$ | −3.0 | 0.00 | 0.95 | 0.65 |
| 4. $MW$ | 158.4 | — | 0.95 | 0.14 |

*Note:* $P - v$ is asymptotically normal, and the right tail is used for rejection. $P - Z_\alpha$ and $P - AEG$ are asymptotically normal, and the left tail is used for rejection. The Maddala–Wu test is distributed chi-squared, and the right tail is used for rejection (Maddala and Wu 1999). The bootstrapped Maddala–Wu test has a nonstandard distribution that must be estimated by simulation; the right tail is used for rejection. The Pedroni tests include aggregate time effects.

regression. A strict interpretation of the results in this article is that an error-correction model is a misspecification and that results from such a model are spurious. A looser interpretation is that, even if prices, income and population are cointegrated, we cannot verify this relationship. Our inability to verify the relationship implies an inability to accurately estimate it. In other words, even if we think a long-run relationship exists, we cannot say with much certainty when prices are in line with fundamentals and when they are not. Forecasts based on our best guess as to the degree of "disequilibrium," such as those from an error-correction model, are therefore highly suspect.

## Conclusion

Many housing market observers have become concerned that house prices have grown too quickly of late and that prices are now too high relative to per capita incomes. Prices will likely stagnate or fall, the argument goes, until they are better aligned with income. This idea is often formalized in the housing literature by asserting a long-run equilibrium relationship between house prices and fundamentals, such as income, population and user cost. The validity of this assumption has important implications for how we model house price dynamics. If the assumption is accurate—so that house prices and fundamentals are cointegrated—then the error-correction specifications of Abraham and Hendershott (1996), Malpezzi (1999) and Capozza *et al.* (2002) are appropriate.

In this article, I have used standard tests to show that there is little evidence for cointegration of house prices and various fundamentals at the national level. I have also shown that bootstrapped versions of more powerful panel-data tests, applied to a panel of 95 U.S. metropolitan areas over 23 years, also do not find evidence for cointegration. This does not mean that fundamentals do not affect house prices, and it does not mean that house prices cannot fall. It does mean that the *level* of house prices does not appear to have a stable long-run equilibrium relationship with the *level* of fundamentals such as income. Thus, the levels regressions found in the literature are likely spurious, and the associated error-correction models may be inappropriate.

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## References

Abraham, J. and P. Hendershott. 1996. Bubbles in Metropolitan Housing Markets. *Journal of Housing Research* 7: 191–207.

Banerjee, A. 1999. Panel Data Unit Roots and Cointegration: An Overview. *Oxford Bulletin of Economics and Statistics* 61: 607–629.

Calem, P. and B. Case. 2005. Co-Movements in U.S. House Price Appreciation Patterns. Working Paper, Federal Reserve Board, Washington, D.C.

Calhoun, C. A. 1996. OFHEO House Price Indexes: HPI Technical Description. Office of Federal Housing Enterprise Oversight (http://www.ofheo.gov/house/download.html).

Capozza, D. R., P. H. Hendershott, C. Mack and C. J. Mayer. 2002. Determinants of Real House Price Dynamics. NBER Working Paper 9262.

Davidson, R. and J.G. MacKinnon. 1993. *Estimation and Inference in Econometrics*. Oxford University Press: New York, NY.

Engle, R.F. and C.W.J. Granger. 1987. Co-integration and Error Correction: Representation, Estimation, and Testing. *Econometrica* 55: 251–276.

Fisher, R.A. 1932. *Statistical Methods for Research Workers*. 12th Edition. Oliver & Boyd: Edinburgh, UK.

Im, K., M.H. Pesaran and Y. Shin. 1997. Testing for Unit Roots in Heterogeneous Panels. Mimeo. Department of Applied Economics, University of Cambridge, MA.

Levin, A. and C.F. Lin. 1992. Unit Root Tests in Panel Data: Asymptotic and Finite Sample Properties. Discussion Paper No. 92–93 (revised 1993), Department of Economics, University of California at San Diego.

Levin, A., C. F. Lin and C.J. Chu. 2002. Unit Root Tests in Panel Data: Asymptotic and Finite-Sample Properties. *Journal of Econometrics* 108: 1–24.

Maddala, G.S. and S. Wu. 1999. A Comparative Study of Unit Root Tests with Panel Data and New Simple Test. *Oxford Bulletin of Economics and Statistics* 61: 631–652.

Malpezzi, S. 1999. A Simple Error Correction Model of House Prices. *Journal of Housing Economics* 8: 27–62.

Meen, G. 2002. The Time-Series Behavior of House Prices: A Transatlantic Divide? *Journal of Housing Economics* 11: 1–23.

Pace, R.K., R. Barry, O. Gilley and C.F. Sirmans. 2000. A Method for Spatial-Temporal Forecasting with an Application to Real Estate Prices. *International Journal of Forecasting* 16: 229–246.

Pace, R.K., R. Barry and C.F. Sirmans. 1998. Spatial Statistics and Real Estate. *The Journal of Real Estate Finance and Economics* 17: 5–13.

Pedroni, P. 1997. Panel Cointegration: Asymptotic and Finite Sample Properties of Pooled Time Series Tests with an Application to the PPP Hypothesis. Mimeo. Department of Economics, Indiana University, Bloomington, IN.

——. 1999. Critical Values for Cointegration Tests in Heterogeneous Panels with Multiple Regressors. *Oxford Bulletin of Economics and Statistics* 61: 653–670.

——. 2001. Panel Cointegration: Asymptotic and Finite Sample Properties of Pooled Time Series Tests with an Application to the PPP Hypothesis. Mimeo. Department of Economics, Indiana University, Bloomington, IN.

Phillips, P.C.B. and S. Ouliaris 1990. Asymptotic Properties of Residual Based Tests for Cointegration. *Econometrica* 58: 165–193.

Poterba, J. 1984. Tax Subsidies to Owner-occupied Housing: An Asset-Market Approach. *Quarterly Journal of Economics* 99: 729–52.

Poterba, J. 1991. House Price Dynamics: The Role of Tax Policy and Demography. *Brookings Papers on Economic Activity* 2: 143–183.

Quah, D. 1990. International Patterns of Growth I: Persistence in Cross-Country Disparities. Working Paper, Massachusetts Institute of Technology, Boston, MA.

Quah, D. 1994. Exploiting Cross Section Variation in Unit Root Inferences in Dynamic Data. *Economic Letters* 44: 9–19.

Reifschneider, D., R. Tetlow and J. Williams. 1999. Aggregate Disturbances, Monetary Policy, and the Macroeconomy: The FRB/US Perspective. *Federal Reserve Bulletin* 85: 1–19.

Shiller, R.J. and P. Perron. 1985. Testing the Random Walk Hypotheis: Power versus Frequency of Observation. *Economic Letters* 18: 381–386.

Topel, R. and S. Rosen. 1988. Housing Investment in the United States. *Journal of Political Economy* 96: 718–740.

## Appendix A: Statistical Definitions

This appendix provides definitions of the panel cointegration tests of Pedroni (1999). Suppose that the hypothesized cointegrating relationship for each city is given by

$$x_{0,i,t} = \alpha_i + \delta_i t + \sum_{m=1}^{M} \beta_{m,i} x_{m,i,t} + e_{it}, \tag{A1}$$

Pedroni suggests the following test statistics:

$$P - v = T^2 N^{\frac{3}{2}} \left( \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \hat{e}_{i,t-1}^2 \right)^{-1}$$

$$P-Z_a = T\sqrt{N}\left(\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{L}_{11i}^{-2}\hat{e}_{i,t-1}^{2}\right)^{-1}\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{L}_{11i}^{-2}(\hat{e}_{i,t-1}\Delta\hat{e}_{it}-\hat{\lambda}_i)$$

$$P-AEG = \left(\tilde{s}_{N,T}^{*2}\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{L}_{11i}^{-2}\hat{e}_{i,t-1}^{*2}\right)^{-\frac{1}{2}}\sum_{i=1}^{N}\sum_{t=1}^{T}\hat{L}_{11i}^{-2}\hat{e}_{i,t}^{*}\Delta\hat{e}_{it}^{*}, \qquad \text{(A2)}$$

where

$$\hat{\lambda}_i = \frac{1}{T}\sum_{j=1}^{k_i}\left(1-\frac{j}{k_i+1}\right)\sum_{t=j+1}^{T}\hat{\mu}_{i,t}\hat{\mu}_{i,t-j}$$

$$\hat{s}_i^2 = \frac{1}{T}\sum_{t=1}^{T}\hat{\mu}_{i,t}^2$$

$$\hat{s}_i^{*2} = \frac{1}{T}\sum_{t=1}^{T}\hat{\mu}_{i,t}^{*2}$$

$$\tilde{s}_{N,T}^{*2} = \frac{1}{N}\sum_{i=1}^{N}\hat{s}_i^{*2}$$

$$\hat{L}_{11i}^{-2} = \frac{1}{T}\sum_{t=1}^{T}\hat{\eta}_{i,t}^2 + \frac{2}{T}\sum_{j=1}^{k_i}\left(1-\frac{j}{k_i+1}\right)\sum_{t=j+1}^{T}\hat{\eta}_{i,t}\hat{\eta}_{i,t-j}$$

and where the residuals $\hat{\mu}_{i,t}$, $\hat{\mu}_{i,t}^{*}$ and $\hat{\eta}_{i,t}$ are from the following regressions:

$$\hat{e}_{i,t} = \hat{\gamma}_i\hat{e}_{i,t-1} + \hat{\mu}_{i,t}$$

$$\hat{e}_{i,t} = \hat{\gamma}_i\hat{e}_{i,t-1} + \sum_{k=1}^{K_i}\hat{\gamma}_{i,k}\Delta\hat{e}_{i,t-k} + \hat{u}_{i,t}^{*}$$

$$\Delta y_{i,t} = \sum_{m=1}^{M}\hat{b}_{mi}\Delta x_{mi,t} + \hat{\eta}_{i,t}.$$

**Appendix B: City Definitions**

| Metropolitan statistical areas: 1978 to 2000 | |
|---|---|
| Akron, OH | Minneapolis-St. Paul, MN-WI |
| Albuquerque, NM | Modesto, CA |
| Ann Arbor, MI | Monmouth-Ocean, NJ |
| Atlanta, GA | Nassau-Suffolk, NY |
| Austin-San Marcos, TX | New Orleans, LA |
| Bakersfield, CA | New York, NY |
| Baltimore, MD | Newark, NJ |
| Baton Rouge, LA | Norfolk-Virginia Beach-Newport News, VA-NC |
| Bergen-Passaic, NJ | Oakland, CA |
| Birmingham, AL | Oklahoma City, OK |
| Boston, MA-NH | Omaha, NE-IA |
| Boulder-Longmont, CO | Orange County, CA |
| Buffalo-Niagara Falls, NY | Orlando, FL |
| Canton-Massillon, OH | Philadelphia, PA-NJ |
| Charlotte-Gastonia-Rock Hill, NC-SC | Phoenix-Mesa, AZ |
| Chicago, IL | Pittsburgh, PA |
| Cincinnati, OH-KY-IN | Portland-Vancouver, OR-WA |
| Cleveland-Lorain-Elyria, OH | Raleigh-Durham-Chapel Hill, NC |
| Columbus, OH | Richmond-Petersburg, VA |
| Dallas, TX | Riverside-San Bernardino, CA |
| Dayton-Springfield, OH | Rockford, IL |
| Denver, CO | Sacramento, CA |
| Des Moines, IA | St. Louis, MO-IL |
| Detroit, MI | Salinas, CA |
| Eugene-Springfield, OR | Salt Lake City-Ogden, UT |
| Flint, MI | San Diego, CA |
| Fort Collins-Loveland, CO | San Francisco, CA |
| Fort Lauderdale, FL | San Jose, CA |
| Fort Wayne, IN | San Luis Obispo-Atascadero-Paso Robles, CA |
| Fort Worth-Arlington, TX | Santa Barbara-Santa Maria-Lompoc, CA |
| Fresno, CA | Santa Cruz-Watsonville, CA |
| Grand Rapids-Muskegon-Holland, MI | Santa Rosa, CA |
| Greensboro-Winston-Salem-High Point, NC | Sarasota-Bradenton, FL |
| Hamilton-Middletown, OH | Seattle-Bellevue-Everett, WA |
| Honolulu, HI | Spokane, WA |
| Houston, TX | Stockton-Lodi, CA |
| Indianapolis, IN | Syracuse, NY |
| Kalamazoo-Battle Creek, MI | Tacoma, WA |
| Kansas City, MO-KS | Tampa-St. Petersburg-Clearwater, FL |
| Lansing-East Lansing, MI | Tucson, AZ |
| Las Vegas, NV-AZ | Tulsa, OK |
| Los Angeles-Long Beach, CA | Vallejo-Fairfield-Napa, CA |

**Appendix B: continued.**

| Metropolitan statistical areas: 1978 to 2000 | |
|---|---|
| Louisville, KY-IN | Ventura, CA |
| Madison, WI | Visalia-Tulare-Porterville, CA |
| Memphis, TN-AR-MS | Washington, DC-MD-VA-WV |
| Miami, FL | West Palm Beach-Boca Raton, FL |
| Middlesex-Somerset-Hunterdon, NJ | Wichita, KS |
| Milwaukee-Waukesha, WI | |