WORKING PAPERS
RESEARCH DEPARTMENT

WORKING PAPER NO. 08-28
RENTS HAVE BEEN RISING, NOT FALLING, IN THE POSTWAR PERIOD

Theodore M. Crone
Swarthmore College

Leonard I. Nakamura
Federal Reserve Bank of Philadelphia

Richard Voith
Econsult Corporation

October 2008
Supersedes Working Paper 06-7
Revision forthcoming in *Review of Economics and Statistics*

RESEARCH DEPARTMENT, FEDERAL RESERVE BANK OF PHILADELPHIA
Ten Independence Mall, Philadelphia, PA 19106-1574 • www.philadelphiafed.org/research-and-data/

RENTS HAVE BEEN RISING, NOT FALLING, IN THE POSTWAR PERIOD

Theodore M. Crone
Swarthmore College

Leonard I. Nakamura
Federal Reserve Bank of Philadelphia

Richard Voith
Econsult Corporation

October 2008

The views expressed here are those of the authors and do not necessarily reflect those of the Federal Reserve Bank of Philadelphia or the Federal Reserve System. We would like to thank Mitchell Berlin, Ronel Elul, Robert Gordon, and David Genesove for detailed comments. We would also like to thank Jack Triplett, Randal Verbrugge, John Krainer, and members of the Brookings Workshop on Economic Measurement and the Federal Reserve System Committee on Applied Microeconomics for valuable comments. David Genesove also kindly supplied the CPI microdata set we have used extensively. This paper is available free of charge at www.philadelphiafed.org/research-and-data/publications/working-papers/.

ABSTRACT

RENTS HAVE BEEN RISING, NOT FALLING, IN THE POSTWAR PERIOD

Until the end of 1977, the U.S. consumer price index for rents tended to omit rent increases when units had a change of tenants or were vacant, biasing inflation estimates downward. Beginning in 1978, the Bureau of Labor Statistics (BLS) implemented a series of methodological changes that reduced this nonresponse bias, but substantial bias remained until 1985. We set up a model of nonresponse bias, parameterize it, and test it using a BLS microdata set for rents. From 1940 to 1985, the official BLS CPI-W price index for tenant rents rose 3.6 percent annually; we argue that it should have risen 5.0 percent annually. Rents in 1940 should be only half as much as their official relative price; this has important consequences for historical measures of rent-house-price ratios and for the growth of real consumption.

JEL Classifications: C81, C82, E31, O47, R21, R31

Correspondence to:
Theodore M. Crone, Swarthmore College, 500 College Avenue, Swarthmore PA 19081, tedcrone@gmail.com (e-mail). Leonard I. Nakamura, Research Department, Federal Reserve Bank of Philadelphia 10 Independence Mall, Philadelphia, PA 19106, 215-574-3804 (office), 215-574-4364 (fax), leonard.nakamura@phil.frb.org (e-mail). Richard P. Voith, Senior Vice President and Principal, Econsult Corporation, 3600 Market Street, Suite 560, Philadelphia, PA 19104, 215-382-1894 (office), 215-382-1895 (fax), voith@econsult.com (e-mail)

RENTS HAVE BEEN RISING, NOT FALLING, IN THE POSTWAR PERIOD

I. Introduction and Overview

This paper constructs a revised estimate of the U.S. consumer price index (CPI) for tenant rents from 1940 to 2000. Until the end of 1977, the U.S. consumer price index for rents tended to omit rent increases when units had a change of tenants or were vacant because the collection method resulted in nonresponses to the survey at these times. This omission biased inflation estimates downward. Beginning in 1978, the Bureau of Labor Statistics (BLS) implemented a series of methodological changes that reduced this nonresponse bias, but substantial bias remained until 1985. Our revised estimate implies that the CPI for rents, which rose by a factor of 4.8 from 1940 to 1985, should have risen by a factor of 10.7. This implies that once we adjust aggregate annual real PCE growth for this mismeasurement, it was 3.5 percent from 1940 to 1985, not 3.7 percent. And we provide evidence that our alternative measures of rental inflation and real output growth are more consistent with other measures of inflation and output growth than the official measures are.

We set up a model of nonresponse bias and parameterize the model from a variety of sources. We then check the parameterization by using a CPI rental microdata set from 1988 to 1992, a period when the biases had been almost entirely corrected and we can directly measure BLS adjustments. The model implies that nonresponse bias from 1942 to 1985 resulted in an understatement of the rental inflation rate of 1.0 percentage point annually from 1942 to 1985. The BLS has estimated that aging bias also affected these data by about 0.4 percentage point annually, so that in total the average annual understatement of rental inflation amounted to 1.4 percentage points annually during this period. From 1940 to 1985, the official BLS CPI-W price index for tenant rents increased 3.6 percent annually. We argue that the true increase was 5.0 percent annually.

Most studies of price mismeasurement have concentrated on upward biases in inflation measures (Boskin et al., 1996; Price Statistics Review Committee, 1961). This paper discusses a case of downward bias in inflation measurement in an important part of the U.S. economy: tenant rents. While one component of nonresponse bias, vacancy nonresponse, has been analyzed in Rivers and Sommers (1983) and corrected by the BLS in 1985, this is the first paper to discuss the nonresponse bias due to loss of tenant contact. Both components of nonresponse bias have disappeared from historical view; neither was mentioned in recent discussions of historical CPI bias by Stewart and Reed (1999) and by Boskin et al. (1996), nor in Moulton's (1997) review of rental inflation biases.[1]

From the mid-1940s forward, researchers at the BLS and in academia suspected that the CPI rental index was downwardly biased (Humes and Schiro, 1948, 1949; Weston, 1972; and Ozanne, 1981). More recently, papers by Crone et al. (2001) and Gordon and vanGoethem (2007) have also suggested such a bias in the historical data. However, the source of much of the bias – whether it was due to response problems, aging bias, or omission of new units – remained murky and has not been corrected in the U.S. Bureau of Economic Analysis' (BEA) deflators for housing services. Our revision implies that real housing services 60 years ago were almost twice as large as BEA estimates. The level of real PCE as a whole in 1942 is about 9 percent higher, and its annual growth rate from 1942 to 1985 is 0.2 percentage point lower; real GDP is 5 percent higher in 1942, and its annual growth rate 0.1 percentage point lower.

The paper most closely related to ours is Gordon and vanGoethem. They estimate bias in the CPI for tenant rents from 1914 to 2003 using data primarily from decennial censuses and the American Housing Survey. For the period 1975 to 2003 they use hedonic methods to estimate both price and quality changes for rental units, including possible vintage effects of advancing technology. Prior to that time they estimate rental price increases by adjusting median rents for quality changes. They argue that between 1935 and 1985, a period they choose to approximate ours, the average downward bias was 1.19 percent annually. Thus, their work is striking confirmation of our results.[2] While their quality-adjusted estimate is somewhat below ours, they find a larger bias estimate, 1.5 percent annually, when they study Evanston, Illinois, using a quasi-repeat rent methodology. Thus, their two estimates of bias bracket ours. Our work goes beyond Gordon and vanGoethem's by providing a rationale for the mismeasurement of rents. We also estimate annual and quarterly rent increases suitable for inclusion in national income accounts.

Our estimates may be important for analyses of housing bubbles in the U.S. One method

[1] The PCE deflator for rent on tenant-occupied nonfarm dwellings has also not been revised; its inflation rate was 3.6 percent for the period from 1940 to 1985, just like the CPI for residential rent.

[2] In an earlier version of this paper, we had estimated the bias to be 1.6 to 1.8 percent, and that is the estimate that Gordon and vanGoethem refer to in discussing our paper.

for estimating equilibrium house prices depends on the stability of the long-run relationship between house prices and rents (see, for example, Gallin, 2004). Our estimate suggests that the BLS measure of rents was double its true value in 1940.

The Bureau of Labor Statistics has long argued that it has been more evenhanded about inflation than its critics have claimed—i.e., its errors have not always resulted in an upward bias in inflation. In this case the BLS removed an important source of bias without a prod from outside critics.

Section II of this paper reviews the history of steps taken by the BLS to correct biases in the CPI rental series. Section III discusses nonresponse bias in the rental CPI and parameterizes the model based on data from the Census Bureau and BLS microdata on rental increases. The parameterized model is used to estimate the bias in rental inflation from 1942 to 1977 and is tested with BLS CPI microdata for 1988-92. Section IV presents our revised rental price index and some additional data on prices and output to suggest that this new estimate is reasonable.

II. History of Changes in BLS Methodology to Correct for Bias in the CPI for Rent

Prior to 1942, nonresponse was not a significant problem in the BLS rental survey because price inspectors obtained their data from the files of real estate agents and large-property owners. The price inspector could directly compare current rents with past rents, even if the tenant had changed. If a unit was vacant, a comparable unit could often be found from the books.

In 1942, the BLS inadvertently created a substantial downward nonresponse bias in its procedure for sampling rents.[3] As wartime rent controls took effect, price inspectors were instructed to obtain rents directly from tenants, which increased the likelihood of nonresponse bias in the rental-price series.[4] Following an initial interview to elicit cooperation and gather data about the unit, the tenant was mailed a rent questionnaire every three months. The nonresponse rate from March to September 1947 was found to be 30 percent—5 percent were returned "unable to locate" and 25 percent were not returned despite follow-up efforts (Humes and

[3] All sample surveys suffer from nonresponse, i.e., incomplete returns from some part of the targeted sample. Pakes (2003), for example, discusses a nonresponse bias in the case of PCs where model exit results in omitting prices that decline, creating an upward bias. In our case, nonresponse bias results in omitting prices that rise.

[4] It was feared that rental increases that evaded or violated rent control laws might not be accurately reported by real estate agents or landlords. These fears were not groundless; Humes and Schiro (1949) report that BLS rents reported twice as many price increases as were authorized in a period in 1947.

Schiro, 1949). When a tenant moved, the mail questionnaire, having been addressed to a previous occupant, would be forwarded or returned to the sender. The BLS rental price inspector would have to ascertain who the new occupant was and solicit his or her cooperation with a new interview and start over again. This process would almost invariably miss the rent increases associated with a change of tenants.

Between 1953 and 1994, the BLS largely corrected nonresponse and other biases in the CPI by taking the following six steps:

(1) 1953: Semiannual rent collection. In 1953, it appears that the rate of rental collection was changed from quarterly to semiannually. Less frequent rental surveys reduced the downward bias in the rental CPI because it reduced the number of data collections occurring when the rent was unchanged, since most rents change only once a year.[5]

(2) 1964: Personal visits and telephone surveys. The mail survey was replaced by personal visit or telephone interview in 1964. It is likely there was some improvement at this time in reaching new tenants, although there is only indirect evidence on this.

(3) 1978: Reducing nonresponse bias and introducing one-month recall bias. Beginning in 1978, a new survey method was introduced: The number of rental units surveyed was reduced and greater effort was expended to obtain higher response rates for the remaining units. Price inspectors could choose to interview the landlord or manager instead of the tenant and typically did so. Response rates from new tenants rose to nearly equal the response rates from continuing tenants.

To speed up the index's sensitivity to recent rent movements, respondents were asked the level of the previous month's rent as well as the current month's rent. The rental index was computed using a composite estimator that included both the one-month change and the six-month change. Defining I(t) as the level of the index at month t, and $R_{t,t-k}$ as the change in rent from k months ago, the rental formula was:

$$I(t) = 0.65\, R_{t,t-1}\, I(t-1) + 0.35\, R_{t,t-6}\, I(t-6). \qquad (1)$$

This formula, known as the composite estimator formula, permitted the CPI measure to reflect current inflation immediately while minimizing noise.[6] Unfortunately, this introduced

[5] Collection of mortgage rate and other price information on the costs of owner-occupied housing was instituted in the 1953 CPI revision, so this was a period in which major changes did occur to the housing index (Lamale, 1956). The 1964 revision announcement confirmed that rent collection had already become semiannual.

[6] That is, the coefficients weighting the six-month change and the one-month change were chosen to minimize the

one-month recall bias because respondents (landlords and managers as well as tenants) often failed to remember increases in rent that had occurred during the previous month. It appears possible that while the BLS does not consider a rent to have increased until the unit is rented, the respondents considered the rent to have increased when the new asking rent was raised, possibly while the unit was vacant.

(4) 1985: Correction of vacancy bias and one-month recall bias. In analyzing CPI rental data in the wake of the 1978 procedural changes, the BLS realized that nonresponse bias remained a problem at vacant units. Vacancy mattered because the BLS treated vacant units as nonresponses, which resulted in a downward bias just like new tenant nonresponse.[7]

Rivers and Sommers (1983) analyzed the BLS sample of rental price increases from late 1979 to early 1981 (Table 1). They divided tenants into continuing tenants, those with six or more months of occupancy, (81.2 percent of the sample) and new tenants (18.8 percent).[8] They noted that rental price increases were much larger when units changed hands. This suggested that some proportion of the true inflation went unmeasured when vacancies were omitted.

They further surmised that if they imputed rents for vacancies and also imputed one-month changes in all vacant units, they could reduce both vacancy bias and one-month recall bias. In their simulations, they used the average new-tenant rental inflation to impute six-month rent increases for vacant units and also used these rates for one-month recall imputations for the vacant units. Doing so eliminated four-fifths of the one-month recall bias.

The BLS decided to impute rents for vacant units using the six-month rent changes for similar units that had turned over for which data were available, beginning in January 1985. Our analysis of the Rivers and Sommers findings implies that correcting the vacancy nonresponse bias alone would have raised the rental inflation rate by 8.7 percent. In addition, the partial correction of one-month recall bias raised the inflation rate by 7.6 percent. Combining these two, introducing the vacancy imputation methodology appears to have raised the measured rental inflation rate by 17.0 percent.[9]

---

decided seasonal patterns that emerge if you use only six-month data, $I(t)=R_{t,t-6}\,I(t-6)$, or only one-month data, $I(t) = R_{t,t-1}\,I(t-1)$.

[7] This is in contrast to the BLS practice for prices other than rents, where transactions are frequent enough so that the BLS feels confident in relying on the asking price, for example, the marked or posted price of a retail item.

[8] The Rivers and Sommers data divide tenants into those with five months or less occupancy and six months or more. It may thus underestimate the proportion of new tenants included in the data, as tenants with more than five months but less than six months occupancy may be in the six months or more category.

[9] This is in line with BLS estimates. In the January 1985 CPI Detailed Report, the BLS stated that the vacancy

(5) 1988: Correction of aging bias. Aging bias refers to the underestimation of rent inflation because of the systematic deterioration in the quality of tenant housing due to aging of the unit. Historically, the BLS has adjusted the change in rent for *observed* quality changes, such as the addition of a room. In 1988 the BLS began adjusting the measure of rental inflation for aging based on the hedonic estimates of Randolph (1988a and 1988b). Correction for aging bias is the only part of this history to which this paper contributes no new analysis.

(6) 1994: Final correction of one-month recall bias. The one-month recall bias introduced in 1978 was completely eliminated in 1994 when the BLS discontinued the use of reported one-month rent increases in estimating rental inflation (Armknecht et al., 1995). At this time, the rent formula was changed so that the monthly rate of rental inflation was calculated as the sixth root of the average six-month inflation rate. The new formula, while free of downward bias, results in roughly a three-month lag in the reporting of changes in the rental inflation rate.

The empirical consequence of one-month recall bias for a sample period was discussed briefly in Armknecht et al. (1995). Rivers and Sommers did not provide an analytical account of the impact of one-month recall bias and the use of the composite estimator. Given the weights in the composite estimator for the six-month and one-month changes in rents, a bias of size e in the recall of the monthly change in rent creates an index bias of 0.2364e. That is, let π be the monthly rate of inflation and $e\pi$ the monthly one-month recall bias. Then if

$$I(t) = 0.65(1+\pi(1-e))\, I(t-1) + 0.35\, (1+\pi)^6\, I(t-6),$$

it is a straightforward use of difference calculus to show that

$$I(t) = (1+\pi(1-0.2364e))^t\, I(0).$$

Rivers and Sommers found that 37 percent of expected one-month changes were omitted, so the expected bias would be 0.37 x 0.2364 = 0.0875. Defining π_t^m as the measured rate of inflation, recall bias thus creates a downward bias equal to

$$\pi_t^m = .9125\pi_t$$

Correcting the nonresponse bias should raise the measured inflation rate by 9.6 percent.[10] The Armknecht et al. (1995) estimate that one-month recall bias was 9 percent of the inflation

imputation adjustment would raise the inflation rate for rents by less than 0.1 percentage point a month. From December 1982 to December 1983, the rental rate rose at an annual rate of 4.8 percent, and from December 1983 to December 1984, at 5.8 percent. Thus 0.1 percent a month would represent 20 to 25 percent of the measured inflation rate, depending on the base against which it was calculated. Vacancy imputation left only a small recall bias, 1.8 percent, to be finally eliminated in 1994.

[10] 1/.9125=1.096.

rate is close to this analytical bias estimate.

III. Modelling and Parameterizing the Consequences of Sampling and Nonresponse

In this section, we set forth a model of the quantitative impact of nonresponse bias. We parameterize the model, using data from a variety of sources, and then we test the parameterization with microdata from the CPI rental survey from 1988 to1992.

Rents in the United States are typically, but by no means always, changed annually when the lease is renewed. More and less frequent adjustment may occur: the lease contract may be for more or less than a year; there may be no lease contract; or the lease contract may provide for rental price changes during its term.[11] But the data indicate that most rent increases occur at roughly annual intervals. This fact influences both how the BLS measures rents and the biases that appear in rental price collection.

III.1 A model of rent collection with nonresponse

Nonresponse bias. In this section, we set up a model to correct the historical CPI for rents for nonresponse bias. The degree of bias will be associated with institutional features of the U.S. economy such as leasing arrangements, dynamic patterns of rent adjustment and turnover, as well as features of BLS price collection. We take the institutional features as given.

The bias correction model assumes that rental units are subject to annual leases. We assume that in a given month at a given rental unit the log rent increases with probability θ (=1/12). When the rent increases, the tenant leaves the unit with probability ρ.

A complicating issue is that the rate of annual inflation at rental units from which tenants move is, on average, higher than at units of continuing tenants.[12] Let us define the rent increase for continuing tenants as π_{Ct}. Where the tenant moves, the rent increase is larger by some fraction, b; for those units, the rent increase is $(1+b)\pi_{Ct}$.[13] Then the rental inflation rate for

[11] The annual lease is the predominant form for rentals. Data from the U.S. Census Bureau's Property Owners and Managers Survey in 1995 (single-family and multifamily units, excluding data not reported or for rent-free units) showed that 44.4 percent of all units had annual leases, 4.0 percent had leases longer than one year, 36.1 percent had leases less than one year, and 15.5 percent had no leases. These facts suggest that while the annual lease is the modal contract under which rental units are occupied, it is by no means universal. Thus, the simple model that underlies our work is an approximation. The survey can be found at www.census.gov/hhes/www/poms.html.

[12] This issue is discussed in Genesove (1999), who argues that landlords and tenants share the ex post surplus of good matches.

[13] An alternative model, which produces equivalent results, would have the rent rise by $(1+b)\pi_{Ct}$ when the tenant

complete data would be $\pi_t = (1+\rho b)\pi_{Ct}$.[14]

Every n months, prices are collected by a BLS price inspector. Nonresponse bias is due to the fact that when the tenant moves, the price inspector is less likely to record a price for the unit, either because the unit stands vacant or because of loss of contact with the tenant. Let us call q_M the probability that a unit where the tenant has moved will have a price recorded, and q_C the probability that a unit with a continuing tenant will be recorded, with $q_M < q_C$. The annualized measured rate of inflation (π_t^m) and the rate of inflation for the complete data (π_t) are then (Appendix 1):

$$\pi_t^m = \frac{1-\rho(1-\frac{q_M}{q_C}(1+b))}{1-n\theta\rho(1-\frac{q_M}{q_C})}\pi_{Ct} \tag{2}$$

$$\pi_t = (1+\rho b)\pi_{Ct} = \frac{(1+\rho b)(1-n\theta\rho(1-\frac{q_M}{q_C}))}{1-\rho(1-\frac{q_M}{q_C}(1+b))}\pi_t^m \tag{3}$$

The correction for nonresponse bias is the coefficient on measured inflation in equation 3. If q_M/q_C is equal to 1, then this coefficient becomes 1, and the measured inflation is the actual inflation rate. There is no nonresponse bias because the bias is due to obtaining fewer observations from units where tenants have moved than units tenants continue to occupy.

All rents omitted when tenant has moved: If rental prices of units are not collected when the tenant at the previous price quotation has moved, then $q_M = 0$ and the equation simplifies to

$$\pi_t^m = \frac{(1-\rho)\pi_{Ct}}{1-n\rho\theta}. \tag{4}$$

If rents are collected annually, $n\theta = 1$, the frequency of sampling would equal the frequency with which prices are changed, and the measured rate of inflation would equal the inflation rate of rents for continuing tenants. Nonresponse bias in that case is only due to the fact that continuing tenants experience lower rates of inflation than new tenants. But with $n\theta < 1$, as the case is with sampling every six months, measured inflation gives too much weight to tenants who are in the

moves, and the next year the new tenant's rent rises by $\pi_{ct+1}/(1+b)$.

[14] We refer to this as the "complete data" rental inflation rate rather than the "true" inflation rate because it is not adjusted for quality biases such as aging.

portion of the annual cycle in which the rent does not increase.

The rate of inflation for the complete data would then be:

$$\pi_t = \frac{(1+\rho b)(1-n\theta\rho)}{1-\rho}\pi_t^m .$$

Modelling vacancy nonresponse. If rental prices are collected when a vacated unit has been reoccupied but not when the vacated unit remains vacant at the time of the next price inspection, we need to calculate the rate of reoccupation. We shall assume a constant monthly rate of reoccupation: for each successive month for a unit whose tenant has left, with probability 1-α a new tenant occupies the unit with a year-long lease at a new fixed price, and with probability α the unit remains vacant. For units occupied in a given month, n months later a price increase will have occurred on average at nθ units; at these units nθ(1-ρ) of the old tenants remain, $\theta\rho(n-\alpha(1-\alpha^n)/(1-\alpha))$ new tenants have moved in, and $\rho\theta\alpha(1-\alpha^n)/(1-\alpha)$ units will have become vacant. To simplify notation, define the ratio of these vacant units to those that experienced a price change as $v \equiv \frac{\rho\alpha(1-\alpha^n)}{n(1-\alpha)}$.

If, for a unit whose original tenant has left, the subsequent rental price is collected when the apartment is reoccupied but not when the unit remains vacant, that is, $\frac{q_M}{q_C} = 1 - v/\rho$, then:

$$\pi_t^m = \frac{(1+\rho b - v(1+b))\pi_{Ct}}{1-v\theta n} \qquad (5)$$

$$\pi_t = (1+\rho b)\pi_{Ct} = \frac{1-v\theta n}{1-v\dfrac{1+b}{1+\rho b}}\pi_t^m .$$

Prior to 1978, the measured rate of inflation followed equation 4, plus aging bias. After 1978, the CPI for rents still suffered from nonresponse due to vacancy and followed equation 5, plus one-month recall bias and aging bias. To examine these relationships quantitatively, we need to estimate the turnover rate (ρ), the vacancy rate (α), the higher rate of inflation experienced by units that turn over (b), and the relative sampling rate of units where tenants move (q_M/q_C).

If we had annual data on each of the parameters of the model for the units in the BLS survey, our measure of nonresponse bias and our corrections to it would be exact. However, we must derive estimates of the parameters of the model from a variety of data sources and will assume that these estimates apply to the BLS surveyed units.[15]

We now turn to estimating the parameters of the nonresponse model.

III.2 Estimates of nonresponse model parameters

Turnover rate, $\rho = 0.344$. The annual turnover rate ρ in our model is the percentage of persons who move *out* of rental units in a given year. There are no published estimates of the turnover rate. The American Housing Survey and the Censuses of Housing both have data on recent movers *into* units. Recent movers into units differ from those who move out of units because they include those who have moved into new (and thus previously unoccupied) rental housing. Annual turnover can be obtained by subtracting new rental units from recent movers. The 1970 Census of Housing provides data on renters who moved into their units between the beginning of 1969 and March 1970.[16] Beginning in 1973, the American Housing Survey (AHS)[17] provides data on renters who moved into their units in the past 12 months. To proxy for the number of renters who moved into new units, we use the number of multifamily (two or more) units constructed during a given year (some new single-family units are rented and some multifamily units are sold for owner occupation, but at least over the period 1970-93, these two have roughly canceled one another out.[18] The estimates are shown, together with the underlying data used in the estimates, in Table 2. For data available from 1970 to 2001, the turnover rate averaged 34.4 percent, varying from 31.1 percent to 37.6 percent, with a standard deviation of 1.86 percentage points. The standard error of the mean is 0.416.

Rental inflation rate adjustments for units where tenants move, b=0.33. Using data from the BLS CPI survey of renters from October 1979 to March 1981, Rivers and Sommers (1983) found that rent increases differed between tenants who had lived in their units less than six

---

[15] Below, we check the reasonableness of this assumption by applying the estimates to a BLS microdata set from 1988 to 1992.

[16] The census period is the previous year and the first three months of the current year. That means that the first quarter is counted twice, a period in which turnover is somewhat lower than during the rest of the year. According to our BLS microdata, 21.6 percent of movers move in during the first quarter of the year; accordingly, we divided this figure by 1.216 to estimate annual movers.

[17] The AHS was known as the Annual Housing Survey from 1973 to 1981, prior to the survey becoming biennial and being renamed the American Housing Survey. We use the new title throughout.

[18] According to the U.S. Census Bureau's American Housing Survey, *Components of Inventory Change, 1980-1993, Pub 8/96*, 95 percent of the multifamily units completed in the same period were rental units. Similar figures apply for 1970 to 1980.

months (new tenants) and those who had lived in their units six months or more (continuing tenants). Among those tenants who had a rent increase, new tenants recorded a six-month increase averaging 11.40 percent (Table 1, occupancy status five months or less, column 5);[19] continuing tenants had an average increase of 8.56 percent (Table 1, six months or more, column 5). Thus, new tenants experienced a 33 percent higher rate of rental inflation when their rents increased.

Rivers and Sommers do not directly provide measures of the error associated with the rent increases. They do provide the results of a multiple-comparison test (Duncan's) that show the differences across means. These enable us to back out upper and lower limits on the standard errors of the means, given that the panel is approximately balanced and assuming that the standard errors for short-term occupants are roughly similar. Using this assumption, the standard error of the mean is between 0.20 and 0.35 percentage point for the groups with occupancy lengths of 1 to 5 months, and between 0.06 and 0.10 percentage point for the group with occupancy greater than 6 months. To compute the standard error for the ratio, we use the Geary-Hinkley transformation, for which we need the means, standard deviations, and correlation between the two terms. The inflation rate of the continuing tenants and those of the new movers is surely positively correlated; it is conservative to then assume a zero correlation. If we do so, then we can calculate a standard error of the ratio b of .044.

Monthly vacancy hazard rates, $\alpha = 0.675$. The parameter α is the probability that a vacant unit is not reoccupied in a given month. This monthly hazard rate is needed to determine the likelihood that an apartment that turns over is vacant when it is surveyed. To estimate the monthly hazard we turn to data on vacancy rates by length of vacancy available in the Housing Vacancy Survey (HVS), which is conducted as part of the Census Bureau's Current Population Survey.[20] The HVS provides information on the proportion of rental vacancies by length of vacancy: units vacant less than six months generally account for 80 percent of units for which the length of vacancy is known. Units vacant less than six months were 5 percent of all rental units from 1970 to 1999 (Table 3). In addition, there are units that are rented but not yet occupied.

[19] Percents have been converted to log percents. This involves some inaccuracy, as average percents and average log percents differ depending on the variance.

[20] Vacancy data are also available from the AHS. The AHS has the drawback that it is conducted from August to November, while the HVS is conducted year-round and is thus unlikely to suffer a strong seasonal bias. The AHS is conducted once every two years, the HVS every month. Samples are roughly the same size; the HVS has about 60,000 units, the AHS about 54,000, but because the HVS units are sampled 24 times in the two-year period during which the AHS is sampled once, the effective size of the HVS sample is much greater.

These appear to be about 1 percent of all units. Assuming that 80 percent of these units have been vacant less than six months, we have total vacancies in a six-month period of 5.8 percent.

Using the model, the one-month vacancy rate is $\rho\alpha\theta$, the total vacancy rate is $\rho\alpha\theta/(1-\alpha)$, and the six-month vacancy rate is $\rho\alpha(1-\alpha^6)\theta/(1-\alpha)$. Assuming that $\rho=0.344$, $\theta = 1/12$, if we set $\alpha = 0.675$, then the percentage of units that are vacant one month or less is 1.94 percent, and the percentage of units that are vacant six months or less is 5.39 percent. This matches the data for 1980-2001 tolerably well.[21] The standard error of the mean for the six-month vacancy rate is 0.164 percent. This implies a standard error for α of 0.0084.

III.3 The sampling rate of units whose tenants have moved, q_M/q_C = 0.2 for 1964-1977.

We do not have direct evidence on the nonresponse rate before 1978. However, we have six months of data that provide some basis for estimating the rate of nonresponse. For the first six months of 1978, the BLS collected CPI data using the old procedures as it was introducing the new one. At the end of the six months, the old procedure showed an inflation rate of 6.4 percent, while the new one estimated 7.0 percent; so the new method shows 8.3 percent *faster* inflation. We can use the measured inflation from the new method, π_t^{MN}, together with parameterizations that we can verify in section V, to estimate π_{Ct} for this six-month period, using equation 5 and our estimate of recall bias. We can then use measured inflation from the old method π_t^{MO} to infer q_M/q_C via equation 2.

Under the new post-1977 methods, vacancy nonresponse bias (equation 5) and recall bias result in: $\pi_t^{MN} = (.9125)\dfrac{1+\rho b - v(1+b)}{1-v\theta n}\pi_{Ct}$, where the coefficient 0.9125 is the adjustment factor for recall bias. Define the coefficient on π_{Ct} as κ. For data from 1977 and earlier, the bias was based on equation 3. We can solve equation 3 for q_M/q_C to obtain:

$$\frac{q_M}{q_C} = \frac{(1-n\theta\rho)\dfrac{\pi_t^{MO}}{\pi_{Ct}} - (1-\rho)}{\rho\left(1+b-n\theta\dfrac{\pi_t^{MO}}{\pi_{Ct}}\right)} = \frac{(1-n\theta\rho)\kappa\dfrac{\pi_t^{MO}}{\pi_t^{MN}} - (1-\rho)}{\rho\left(1=+b-n\theta\kappa\dfrac{\pi_t^{MO}}{\pi_t^{MN}}\right)},$$

Here we can substitute in π_t^{MO}/π_t^{MN} to infer q_M/q_C. We have only six observations on

[21] However, this model does not match the data well beyond six months. The reoccupation rate tends to fall over time; indeed, the vacancy rate in the simple model falls too steeply to match the data from two to four months to four to six months; so it should be kept in mind that α has been calibrated to fit the average three-month and six-month vacancy rates. In experiments with the model where n changes, the model has a low vacancy rate when n=12.

π_t^{MO} and π_t^{MN}. The estimated ratio of π_t^{MO}/π_t^{MN} is 0.9121; the corresponding estimate of q_M/q_C is 0.163. The bootstrap standard error is 0.0677. If we construct an 80 percent confidence interval, the upper bound is 0.969; 90 percent of the probability mass lies below this estimate of π_t^{MO}/π_t^{MN}, for which the corresponding estimate of q_M/q_C is 0.397.

In appendix 2, we discuss two additional issues: whether these estimates might be biased because the new method imparts a different seasonality or timing to the data. We show that our estimates of q_M/q_C are unaffected by these issues or even that our estimate might be too high. Overall, then, we have a reasonable argument that q_M/q_C was 0.2 or less in the period before 1978. We shall assume this 0.2 rate from 1964 to 1977.

For the period before 1964, given the high rate of nonresponse, it seems reasonable that q_M =0; that is, no new tenants were sampled when surveys were conducted by mail. The sampling rate for new tenants probably increased after 1964 when personal visits were instituted.

Bias correction factors. In Table 4 we summarize the correction factors used to construct our new index of rental price inflation. It gives a chronology of the BLS' changes in its rental collection methods and our model estimates for the impact of each change. For the entire period of 1942 to 1977 we use the model parameters: higher rate of rental increase for new tenants, b = 0.33; turnover rate, ρ = 0.344; years in a month, θ = 1/12; and vacancy hazard, α = 0.675.

Before 1942, our arguments suggest that the BLS methodology was biased only because of the omission of an aging bias correction; this we call method 1. From 1942 to 1952, the CPI's nonresponse bias was unusually large because of quarterly data collection and aging bias; we call this method 2. From 1953 to 1963, the mail survey continued to result in very few rent collections from units that changed hands under semiannual collection (method 3) and the BLS inflation measure must be revised upward by 40.5 percent in addition to correcting for aging bias. From 1964 to 1977 (method 4), the telephone survey raised response rates, and the nonresponse bias implies an upward revision of 28.5 percent. From 1978 to 1985 (method 5), when managers and landlords could be contacted and the price inspectors' contact with units rose substantially, vacancy bias and one-month recall bias together resulted in a bias factor of 1.181. Beginning in 1985 (method 6), vacancy imputation eliminated nonresponse bias, and only a small amount of one-month recall bias remained in addition to aging bias. Aging bias was corrected in 1988, at which point only a portion of the one-month recall bias remained (method 7). Beginning in January 1994, when the composite estimator was abandoned, the CPI rental index required no adjustment (method 0).

The bias adjustment factors depend crucially on the model parameters. To provide some measures of how much these factors might fall if the model parameters are misestimated, we can construct bias factors that correspond to the 10 percent lower bounds for each of the parameters. For parameter b, the 10 percent lower bound is .274 instead of .33. Substituting this value would reduce the method 4 estimate of 28.5 percent to 26.9 percent. For parameter ρ, the 10 percent lower bound is .339, instead of .344, this would reduce the method 4 estimate from 28.5 percent to 28.2 percent. For parameter q_M/q_C, the 10 percent lower bound is .397, instead of 0.20. This would reduce the method 4 estimate from 28.5 percent to 19.2 percent. If we include all three of these lower bounds together, the method 4 estimate is reduced from 28.5 to 18.6 percent. The parameter α is used only in the method 5 estimate. For this parameter, the 10 percent lower bound is 0.664 instead of 0.675. This would reduce the method 5 estimate from 18.1 percent to 17.7 percent.

Thus the main parameter estimate that could sharply change the bias adjustment factors is the parameter q_M/q_C. It is for this reason that we have focused on the estimation of this parameter and its possible biases.

III.4 Testing the model of nonresponse bias: Simulation with BLS microdata

In this section, we test the validity of our parameterized model by using the CPI microdata for rents for the period January 1988 to December 1992. In this period, the BLS was still collecting information from renters about the previous month's rent and the current month's rent and using the composite estimator; it imputed missing data for vacancies and other nonresponding units; and it adjusted the data for aging bias. The data set includes information on each housing unit sampled by the BLS. For each unit and collection period, the data set has information on the length of occupancy (one to six months and more than six months); the type of structure; the completeness of the interview or a reason for failure to obtain information; the current month's rent—either actual or imputed by the BLS; and last month's rent, actual or imputed. The data also provide information on which observations have been imputed and whether the tenant is continuing from the last rent observation or a new occupant. It is thus a very good data set for verifying whether the data the BLS actually used conform to our model behavior, since it provides us the data necessary to compute the impact of changes in BLS practices.[22] However, it does not provide information on the nonresponse rate prior to 1978.

Rental inflation estimates based on microdata. Table 5 shows the microdata estimates of alternative BLS methods of data collection. The first number reported for each method shows what the measured inflation rate would be, using only the six-month changes in the microdata; so these data omit any one-month recall bias. We carry over the method numbering from Table 4. Method 0 includes the imputed rents for vacant units in the microdata and represents the current methodology except it does not include aging bias. It thus represents complete data. Method 3 excludes all recent movers (corresponding to the procedures used from 1953 to 1964, while method 4 excludes 80 percent of recent movers (corresponding to the procedures used from 1965 to 1977, using our estimate of $q_M/q_C = 0.2$). Method 5 includes only the rent data actually collected from respondents, mimicking the method from 1978 to 1984, with vacancy bias. Method 6 is the method used during the period from 1985 to 1993, ex aging bias, complete data with a small one-month recall bias. The second number reported for methods 5, 6, and 0 shows the one-month data. The third number reported for each of these three methods shows the six-month and one-month data combined using the one-month estimator. Table 6 shows the data from Table 5 in ratio form, enabling us to compare the ratios implied by our model (as shown in Table 4) to the ratios from the microdata-based simulation.

Bias correction factors. In Table 4 we presented the correction factors that our parameterized model suggests for different periods as the BLS changed its rental collection and processing methods. We first discuss how these compare to simulated data. (We are unable to duplicate method 2, the period of quarterly collection from 1942 to 1952, because in the period from which the microdata are taken, there was only semiannual collection.) Note that the parameterization of our model uses no data from the microdata set.

From 1953 to 1963, our model suggests that inclusion of new tenants and vacancies would raise the measured inflation rate by 40.5 percent. In the microdata, inclusion of new tenants raises the measured inflation rate by 39.5 percent. From 1964 to 1977, our revision

[22] The data set does not have the weights the BLS used to blow up the sample observations to the universe. A simulation using BLS methodology at the time reveals a very small difference in the official non-seasonally-adjusted rental inflation and the simulated rental inflation using our unweighted data: our simulation estimates rental inflation of 3.461 percent (not seasonally adjusted annual rate, in logs) from June 1988 to December 1992, compared to 3.438 percent in the published data. The difference is reduced even further if we avoid the problems of seasonality by using the annual averages for 1989 and 1992 (the difference in inflation rates over the period is just 0.003 percentage point: 3.369 percent annually in our simulation to 3.363 percent in published data). Throughout these simulations we will use data that average the full year 1992 and the full year 1989.

suggests that the correction factor in this period needed to eliminate nonresponse bias was 28.5 percent, somewhat below the simulation ratio of 32.7 percent. From 1978 to 1984, the correction factor from the vacancy nonresponse bias model is 8.6 percent, while the simulation data suggest an 11 percent upward correction.

All errors in our ratios to the complete data were less than 25 percent and in most cases much less. All the larger errors imply that our correction factors are too conservative. Note further that the Rivers and Sommers data we used to calibrate the model were from a period of close to double-digit inflation, while in the simulation period inflation was about 3 percent. Thus, it appears likely that our formulas are almost certainly a better approximation to the true inflation rate than the original published data.

The composite estimator and one-month recall bias in microdata. Now let us test our one-month recall bias formula using the 1988-1992 rent microdata. During this period, the BLS was using imputations to fill in data for a large proportion of observations. It used the six-month relatives for recent movers to impute the six-month relatives for vacancies and other nonresponders and obtain the current rental price. It also imputed estimates of one-month inflation rates by assuming a proportion of the six-month rate was appropriate.

In the actual one-month rental increase data (without vacancy imputations), the average annualized rate of increase (1.676 percent) is only 60.5 percent as much as the average annualized rate of increase in the six-month actual changes (2.767 percent). The vacancy imputations raise the annualized rate of increase in the six-month rental increase data by 11 percent, from 2.767 percent to 3.071 percent, and they raise the annualized rate of increase in the one-month rental increase data by 69 percent, from 1.676 to 2.835 percent.

Using the methods corresponding to the 1978 to 1984 period, the one-month recall bias in the composite estimator reduced the measured inflation rate from 2.767 percent to 2.509 percent. To eliminate this one-month recall bias thus raises the rental inflation rate by 10.3 percent, close to our modeled estimate of 9.6 percent.

On the other hand, vacancy imputations not only correct the six-month data; they also correct the one-month data. As a result, the impact of the composite estimator on the imputed data is to lower the measured inflation rate by only 2 percent, from 3.071 to 3.010 percent. This implies that e has been reduced to about 0.08 π. This closely matches Rivers and Sommers's expected impact of vacancy imputations on recall bias.

The value of the composite estimator in smoothing the six-month relatives is evident in Figure 1. Here we have graphed separately the two parts of the composite estimator, using the one-month price changes to create two inflation series that use only the one-month data (using the formula $I_1(t) = R_{t,t-1} I_1(t-1)$) and the six-month price changes to create three inflation series (using the formula $I_6(t) = R_{t,t-6} I_6(t-6)$.) For the one-month data, we create a series without the vacancy imputations (actual data) and with the imputations (complete data). The strikingly faster rate of inflation with the imputations can be seen clearly. For the six-month data, we create series corresponding to continuing tenants only (the 1953 to 1963 method), all tenants without vacancy imputations (actual data), and all tenants including vacancy imputations (complete data).

As can be seen, in all cases the one-month relatives are distinctly smoother than the six-month relatives. The composite estimator gives a far more stable series as well, precisely what the formula was designed to do. It is tempting to use a formula such as $I_6(t) = R_{t,t-6} I_6(t-6)$, because this formula does not create an artificial three-month lag in the published series as the current BLS formula does. But it introduces noise into the series in the form of a substantial sawtooth.

IV. A New Measure of Rental Inflation, 1940-2001

In 1999 Stewart and Reed published an adjusted CPI that incorporated the adjustments for one-month recall bias and aging bias into the historical rental inflation series. We believe that to correctly adjust the historical data, a further adjustment needs to be made for nonresponse bias. In creating our new estimates of rental inflation, we developed estimates of the impact of increased response rates for new renters, one-month recall bias, and vacancy imputation and have used estimates of aging bias from the BLS. Our new rental price series imply that historical measures of U.S. aggregate inflation, including the personal consumption expenditure (PCE) deflator, the CPI, and the CPI-U-X1, included a downward bias in rents of 1.4 percentage points a year over the entire period from 1940 to 1985.

Annual rental price indexes for December of each year from 1940 to 2000 for our revised estimates of the rent series are presented in Table 7.

IV.1 Comparing alternative rental inflation estimates

In this section we assess the reasonableness of our revised CPI for rents by comparisons with a number of other data series. Does our new series appear to be more closely aligned with median rents and other data series on inflation and real growth?

Table 8 shows the relationship between median gross rent and rental inflation data. As the final column shows, the revision reduces the gap between the CPI rental inflation and the median rent growth rate, but in the period 1940 to 1985 does not eliminate it. From 1985 to1995, however, our revised rental inflation was only roughly 0.2 percentage point less than median rent annually, which implies a small quality increase over the period. In the most recent period, 1995 to 2001, we do not revise the CPI rent measure, as we believe that tenant rents were correctly calculated. In this period, the rental inflation measure grew 0.3 percentage point faster than median rent, implying that the quality of the rental stock was falling modestly.

Table 9 gives old and new estimates of rental inflation from 1975 QIV to 2001 QIV together with econometric estimates of rental inflation based on microdata from the American Housing Survey. These econometric estimates are based on Crone et al. (2001); we use fourth quarter data to match the timing of the American Housing Survey. The rental inflation measures are based on Box-Cox hedonic regressions and on repeat rent models. The Box-Cox rental inflation rates are relatively close to those of the adjusted CPI for rent, particularly in the period from 1975 to 1985 when the CPI adjustments are the largest. These provide some supportive evidence for the reasonableness of the adjustment.

On the other hand, the repeat rent estimates that use the panel subsamples of the AHS are closer to the unadjusted CPI rent measures. One difference is that the repeat rent measures do not include an adjustment for aging bias. However, that accounts for only 0.4 percentage point of the 2.0-percentage-point gap between the two series during the crucial period from 1975 to 1983. A more important issue is that the repeat rent estimates may suffer from nonresponse bias, since a high proportion of observations are missing in the panel data.

Table 10 shows long-term annual inflation rates for the periods 1940 to 1985 and 1985 to 2001. The PCE tenant rent and owner-occupied rental equivalent housing services price indexes closely mirror the long-run inflation rate of the CPI for tenant rents of the BLS, since the BEA depends primarily on the CPI for tenant rents in constructing these deflators. In the period before 1985, these official rent estimates tend to be well below our revised rent estimate, the median gross rent, and the BEA's residential fixed investment chain price deflator. The official rent inflation estimates are also well below all the other U.S. aggregate price inflation measures. We use the CPI-W excluding shelter because that provides a well-known measure of the CPI that excludes rents (it also excludes the problems associated with the use of the mortgage interest rate in the CPI before 1983). We also include the personal consumption deflator, the GDP deflator, and the PPI all-items price index (linked to the old wholesale price index). These data all suggest that the published rental inflation rates in this period are anomalously low.

In sharp contrast, in the period from 1985 to 2001, where we have argued that the official rent inflation measures are generally correct, all the rental inflation measures are generally rising faster than the aggregate price measures, consistent with trend productivity growth in construction being slower than in other sectors of the economy.

Comparing the two periods, the unrevised CPI and PCE rental inflation measures show almost no deceleration between the two periods, slowing by less than 0.2 percentage point. This lack of deceleration stands in contrast to alternative measures of inflation that show deceleration of between 1.6 and 3.6 percent. The revised CPI rental measure shows a deceleration much closer to the other price measures. This also suggests that the unrevised measures are anomalous, unless there is a sharp break in trend productivity growth for housing.

Table 11 compares the growth rates of the two PCE measures of housing services with alternative measures of real activity. The revised measure of real PCE housing services is constructed by deflating owner-occupied, tenant, and farm dwellings with the revised CPI-W. (Other – primarily hotels – is small and left unchanged.) The BEA net stock quantity index for residential fixed assets is constructed by the perpetual inventory method and reflects the real stock of housing net of depreciation. If there is a sharp break in construction productivity, which might drive the anomalous movements discussed above, then one would expect a relatively stable relationship between the BEA's measure of the residential net stock and PCE for real housing services, since the housing services are those provided by the stock of housing.

From 1985 to 2001, the BEA's measure of housing services grows at the same rate as the net stock, as one would expect. However, from 1940 to 1985, the BEA's measure grows much faster, consistent with the possibility that inflation has been understated and housing services growth overstated. By contrast, the relationship between our revised measure of housing services and the net stock is relatively stable. The revised measure of housing services and the net stock measure–though derived from entirely different procedures and data–tell a broadly consistent story, while the unrevised measure does not.

Table 11 further shows that the BEA's measure of housing services grew faster from

1940 to 1985 than the rate of residential fixed investment, real gross domestic product, and real personal consumption expenditures. By contrast, from 1985 to 2001 it grew either about as fast as or slower than other measures of real activity. The last two rows show that both payroll and population growth decelerated over the two periods, in line with the deceleration of other measures of economic activity. These data are also supportive of the revised estimates of housing services growth and thus of the revised CPI rental inflation measures.

Rent-price indexes. Unfortunately, a constant-quality house price index going back to 1940 does not exist. The constant-quality index with the longest time series, Freddie Mac's Conventional Mortgage Home Price Index (CMHPI), goes back to 1970. We can thus study rent/price ratios using our data and comparing it to the original CPI-rent series, from 1970 to 2001, as quarterly series. We would expect such a ratio to be mean reverting, rather than possessing a unit root.

Using the ADF-GLS test of Elliott, Rothenberg, and Stock (1996), we find that with our rental series in the numerator of the rent/price ratio, we can reject the hypothesis of a unit root at the 5 percent level, setting lag lengths with SIC. On the other hand, with the CPI rental series we cannot reject the hypothesis of a unit root at the 10 percent level. Using alternative criteria for setting lag lengths (AIC, HQ, MAIC, MSIC, MHQ), with our series we always reject at least at the 10 percent level, and with the BLS series we cannot reject at the 10 percent level. As a further check, we create a hybrid house price index with the Office of Federal Housing Enterprise Oversight housing price index since 1975Q1 and the CMHPI from 1970 to 1975Q1. We obtain similar results: we can reject a unit root with our rent series and cannot reject with the BLS rental series.

We believe this is reasonable evidence that our rental price series will be more useful in studying rent-price dynamics than the existing BLS CPI for rents.

V. Summary

We have argued in this paper that the rate of rental inflation was quite substantially underestimated in the period from 1942 to 1985, by about 1.4 percentage points annually. The BLS long suspected a problem with the data and fixed the bias, step by step, over the course of decades. In this paper, we have modeled the impact of nonresponse bias–the main source of the rental inflation bias–and calibrated our model with data from the American Housing Survey, the

Housing Vacancy Survey, and a BLS microdata study from the period 1979 to 1981. We then verified our estimates using BLS microdata from the period 1988 to 1992. Finally, we have shown that our estimates of substantial bias are consistent with other economic statistics, using a variety of alternative measures of inflation and growth.

References

Armknecht, Paul A., Brent R. Moulton, and Kenneth J. Stewart, "Improvements to the Food at Home, Shelter, and Prescription Drug Indexes in the U.S. Consumer Price Index," BLS Working Paper 263, 1995.

Boskin, Michael J., E. Dulberger, R. Gordon, Z. Griliches, and D. Jorgenson, "Toward a More Accurate Measure of the Cost of Living," Final Report to the Senate Finance Committee, December 4, 1996.

Bureau of Labor Statistics, *Consumers' Prices in the United States 1942-48: Analysis of Changes in the Cost of Living,* Bulletin 966 (1949).

Bureau of Labor Statistics, *The Consumer Price Index: History and Techniques*. (May 1966).

Bureau of Labor Statistics, *Consumer Prices in the United States, 1959-68: Trends and Indexes*. Bulletin 1647 (1970).

Bureau of Labor Statistics, *The Consumer Price Index: Concepts and Content Over the Years*. (May 1978).

Crone, Theodore M., Leonard I. Nakamura, and Richard Voith, "Measuring American Rents: A Revisionist History," Federal Reserve Bank of Philadelphia Working Paper 01-8, July 2001.

Elliott, Graham, Thomas J. Rothenberg, and James H. Stock, "Efficient Tests for an Autoregressive Unit Root," *Econometrica* 64, July 1996, 813-36.

Gallin, Joshua, "The Long-Run Relationship between House Prices and Rents," Finance and Economics Discussion Series 2004-50, Board of Governors of the Federal Reserve System, 2004.

Genesove, David, "The Nominal Rigidity of Apartment Rents," NBER Working Paper 7137, May 1999.

Gordon, Robert J., and Todd vanGoethem, "Downward Bias in the Most Important CPI Component: The Case of Rental Shelter, 1914-2003," in Ernst R. Berndt and Charles M. Hulten, eds., *Hard-to-Measure Goods and Services, Essays in Memory of Zvi Griliches*, University of Chicago Press, Chicago, 2007.

Humes, Helen, and Bruno Schiro, "The Rent Index--Part 1, Concept and Measurement," *Monthly Labor Review* 67 (December 1948), 631-637.

Humes, Helen, and Bruno Schiro, "The Rent Index--Part 2, Methodology of Measurement," *Monthly Labor Review* 68 (January 1949), 60-68.

Lamale, Helen Humes, “Housing Cost in the Consumer Price Index,” *Monthly Labor Review* 79 (February 1956), 189-196, and (April 1956), 442-446.

Moulton, Brent R., “Issues in Measuring Price Changes for Rent of Shelter,” Paper presented at Conference on Service Sector Productivity and the Productivity Paradox, April 1997.

Ozanne, L., “Expanding and Improving the CPI Rent Component,” in J. Tuccillo and K. Villani, eds., *House Prices and Inflation*. Urban Institute Press, Washington, DC: 1981, 109-122.

Pakes, Ariel, “A Reconsideration of Hedonic Price Indexes with an Application to PC’s,” *American Economic Review* 93 (December 2003), 1578-96.

Price Statistics Review Committee, *The Price Statistics of the Federal Government*. NBER, New York, 1961.

Randolph, William C., “Estimation of Housing Depreciation: Short-Term Quality Change and Long-Term Vintage Effects,” *Journal of Urban Economics* 23 (1988a), 162-178.

Randolph, William C., “Housing Depreciation and Aging Bias in the Consumer Price Index,” *Journal of Business and Economic Statistics* 6 (1988b), 359-371.

Rivers, Joseph D., and John P. Sommers, “Vacancy Imputation Methodology for Rents in the CPI,” *Proceedings of the ASA Economics and Business Section*, 1983, 201-205.

Stewart, Kenneth J., and Stephen B. Reed, “Consumer Price Index Research Series Using Current Methods, 1978-1998,” *Monthly Labor Review* 122 (June 1999), 29-38.

Weston, Rafael Rom, *The Quality of Housing in the United States*, Ph.D. Thesis, Department of Economics, Harvard University, September 1972.

Appendix 1

Calculation of Rental Inflation Adjustments for Nonresponse Bias

Assumptions about parameters in model			
Event	Probability of event	Log change in rental	Probability of measurement
Lease in force	$1-n\theta$	0	q_C
Lease ends, tenant stays	$n\theta(1-\rho)$	π_{Ct}	q_C
Lease ends, tenant leaves	$n\theta\rho$	$(1+b)\pi_{Ct}$	q_M

Quantity of successfully recorded responses per measurement attempt: $q_C(1-n\theta\rho)+q_M(n\theta\rho)$

Measured inflation per measurement attempt: $q_C(n\theta(1-\rho)\pi_{Ct}) + q_M(n\theta\rho(1+b)\pi_{Ct})$

Define the annualized inflation rate as π^m_t

Measured inflation for time period $n\theta$:

$$n\theta\pi_t^m = \frac{q_C n\theta(1-\rho)\pi_{Ct} + q_M n\theta\rho(1+b)\pi_{Ct}}{q_C(1-n\theta\rho)+q_M(n\theta\rho)}$$

which simplifies to:

$$\pi_t^m = \frac{1-\rho\left(1-(\frac{q_M}{q_C}(1+b))\right)}{1-n\theta\rho(1-\frac{q_M}{q_C})}\pi_{Ct}$$

and also:

$$\frac{q_M}{q_C} = \frac{(1-n\theta\rho)\pi_t^m-(1-\rho)\pi_{Ct}}{\rho\left((1+b)\pi_{Ct}-n\theta\pi_t^m\right)}$$

Appendix 2

Seasonality and rental inflation acceleration and the new and old method comparison, December 1977 to January 1978

Seasonality. Did the new method impart a new seasonal pattern to the data? If so, then because our data cover only a six-month period, there might be a false impression imparted by the unadjusted data. So we used the U.S. Census Bureau's X-12 seasonal adjustment program as implemented in Eviews 5.0, first on the old method data, from January 1954 to June 1978, and then on the new method data, from December 1977 to December 2004. The resulting data are shown in Appendix Table 2. They show an even faster rate of acceleration. The estimated ratio of π_t^{MO}/π_t^{MN} is 0.7972, with a bootstrap standard error of 0.09662 and an 80 percent confidence interval from 0.6734 to 0.9210. The central tendency is implausibly low, implying a negative value of q_M/q_C; the upper limit of the confidence interval implies a q_M/q_C of 0.186. The seasonal adjustment in this case raises the standard deviation of the log changes in the data rather than smoothing them out, suggesting that the seasonal adjustment may be inaccurate; seasonal adjustment is generally least accurate at the beginning and the end of series.

Timing and inflation acceleration. Is there an issue with timing? The composite estimator was designed to provide more current data than an index without one-month recall. Although we know that the trend impact of one-month recall bias would be to reduce the measured inflation rate, some of this impact might be offset if rental inflation were accelerating, in which case the more timely indicator would show more inflation. This does not appear to be the case. The measured not seasonally adjusted inflation rate, annualized, for the six months from December 1977 to June 1978 was 6.85 percent, while for the three months from June 1978 to August 1978 it was 6.88 percent.

Data for December 1977 to June 1978				
	Not Seasonally Adjusted		Seasonally Adjusted, X-12	
	Old Method	New Method	Old Method	New Method
December 1977	66.9	66.9	66.9	66.7
January 1978	67.2	67.3	67.1	67.2
February 1978	67.6	67.6	67.5	67.6
March 1978	68.0	68.0	67.9	68.1
April 1978	68.4	68.4	68.3	68.7
May 1978	68.7	68.9	68.7	69.0
June 1978	69.0	69.2	69.0	69.3

TABLE 1 **SIX-MONTH RENT INCREASES FOR CONTINUING AND NEW TENANTS** Data collected October 1979 to March 1981, reflecting changes from April 1979 to March 1981						
(1) Occupancy status	(2) Number surveyed	(3) Number with six-month rent change	(4) Proportion with rent change	(5)* Average rent change for units with change	(6)* Average rent change for all units	(7)* Average rent change for all units, annualized
6 months or more	37144	17243	.464	8.56	4.07	8.1
5 months or less	8614	6939	.806	11.40	9.28	18.6
all occupants	45758	24182	.528	9.38	5.07	10.1
vacancies	3833					
other nonresponses**	3868					

Data from Rivers and Sommers, 1983, pp. 202-203, tables "Analysis of Six-Month Rent Changes by Length of Occupancy" and "Interview Classification."
*Log percent changes
**Includes no one at home (2619), refusal (745), and other (504).

TABLE 2 ESTIMATE OF RENTAL TURNOVER RATE				
	(1) Vacancy Survey	(2) AHS & Census	(3) Housing Completions	(4) Estimated Turnover Rate
Year	Occupied rental units	recent movers	Multifamily	=[(2)-(3)]/(1)
1970	22806	7707*	618.0	31.1%
1971	23266		688.1	
1972	23849		839.9	
1973	24425	8892	902.3	32.7%
1974	24943	9426	792.7	34.6%
1975	25462	9698	445.9	36.3%
1976	25897	9924	341.7	37.0%
1977	26324	10302	397.0	37.6%
1978	26810	9940	496.3	35.2%
1979	27174	9885	570.6	34.3%
1980	27416	10116	547.0	34.9%
1981	28709	10862	446.5	36.3%
1982	29495		373.6	
1983	29894	9958	464.9	31.8%
1984	30675		623.6	
1985	31736	12166	632.0	36.3%
1986	32302		638.3	
1987	32602	12275	548.3	36.0%
1988	33292		446.0	
1989	33734	12303	397.5	35.3%
1990	33976		343.3	
1991	34242	12230	254.8	35.0%
1992	34568		193.4	
1993	35184	11524	153.2	32.3%
1994	35557		185.0	
1995	35246	12251	246.5	34.1%
1996	34943		283.0	
1997	35059	11969	284.6	33.3%
1998	34896		315.4	
1999	34830	11349	333.3	31.6%
2000	34470		332.7	
2001	34417	11641	314.7	32.9%
2002	34826		321.4	
average				34.4%

Sources: (1) Housing Vacancy Survey, (2)*Census of Housing (1970), 9372 divided by 1.216 to account for 5 quarter period (see text), American Housing Survey (1973-2001), (3) Residential Construction Survey.

TABLE 3 RENTAL VACANCY RATES BY LENGTH OF VACANCY			
	1960, 70, 75*	1980-2001*	Model estimates**
Total vacancy	6.47	7.23	5.95
1 month or less	2.14	2.20	1.94
1 to 2 months	0.95	1.27	1.31
2 to 4 months	1.08	1.36	1.48
4 to 6 months	0.58	0.74	0.67
less than 6 months	4.75	5.58	5.39
6 months or more	1.73	1.66	0.56

Source: Census Bureau. *Housing Vacancy Survey*.
* Data are averages of available data. Dates published are 1960, 1970, 1975 and 1980 to 2001.
** Model uses formula for cumulative vacancy rate: $\frac{\rho\alpha\theta(1-\alpha^n)}{1-\alpha}$ where n is the number of months vacant, with ρ=.344 and α=.675.

TABLE 4 **CORRECTIONS FOR BLS RENTAL INCREASES** (Model estimates of the multiplicative factor needed to adjust CPI to true rental inflation rate given various parameter estimates)				
Method	Periods	Problems	Parameters	Formulas to create revised inflation rate
			All rows: θ=1/12, b= .33 ρ=.344	
1*	Before January 1942	Aging bias		π_{BLS1} + .36
2*	January 1942 to December 1952	Response bias, quarterly collection, aging bias	q_M/q_C = 0, n= 3,	1.551 π_{BLS2} +.36 %
3*	January 1953 to December 1963	Response bias, semiannual collection, aging bias	q_M/q_C =0, n=6	1.405 π_{BLS3} + .36 %
4*	January 1964 to December 1977	Response bias, semiannual collection, aging bias	q_M/q_C = 0.2, n=6	1.285 π_{BLS4} + .36 %
5**	January 1978 to December 1984	Vacancy bias, one-month recall bias***, aging bias	n = 6, α=0.675	1.190 π_{BLS5} + .36 %
	of which:	Vacancy bias	n = 6, α = 0.675	1.0859
		One-month recall bias***	d = 0.2364 e = 0.37 π	1.0959
6	January 1985 to December 1987	One-month recall bias (1/5 remaining)***, aging bias	d = 0.2364 e = 0.074 π	1.018 π_{BLS6} + .36 %
7	January 1988 to December 1993	One-month recall bias (1/5 remaining)***		1.018 π_{BLS7}
0	January1994 to present			π_{BLS0}

*Turnovers partially omitted formula: $$\frac{(1+\rho b)(1-n\theta\rho(1-\frac{q_M}{q_C}))}{1-\rho(1-\frac{q_M}{q_C}(1+b))}$$ (1942-1977)

**Vacancies omitted formula: $$\frac{1-\theta v n}{1-v\frac{1+b}{1+\rho b}} \quad \text{where } v=\frac{\rho\alpha\left(1-\alpha^n\right)}{n(1-\alpha)}$$ (1978 to 1984)

***One-month recall bias: I(t) = (1+μ-de) I(t-1)

TABLE 5 **SIMULATION OF ALTERNATIVE RENT METHODOLOGIES:** **AVERAGE ANNUALIZED INFLATION RATES 1989 TO 1992 (LOG PERCENT)** (no adjustments for aging bias applied)		
Method applied from Table 4 (vintage to which method is applied)	Corresponding subset of microdata used	Average annualized inflation rate
Method 3 (1953 to 1963)	Excludes all recent movers, no imputations 6-month changes only	 2.201
Method 4 (1964 to 1977)	Excludes 80 percent of recent movers (q_M/q_C=0.2), no imputations 6-month changes only	 2.314
Method 5 (1978 to 1984)	All data except imputations 6-month changes only 1-month changes only (without correcting for one-month recall bias) Weighted average	 2.767 1.676 2.509
Method 6 (1985 to 1993) Method 0 (1994 to present)	All data, including imputations 6-month changes only 1-month changes only (without correcting for one-month recall bias) Weighted average	 3.071 2.835 3.010

Source: BLS microdata; see text.

TABLE 6 COMPARISON OF IMPLIED CORRECTION FACTORS FROM SIMULATIONS BASED ON BLS MICRODATA, 1988 TO 1992, AND PARAMETERIZED MODEL ESTIMATES			
	Methods used to calculate ratios for correction factors implied by simulations	Correction factors implied by simulations	Correction factors implied by parameterized model estimates
1953 to 1963	method 0 (without correction for aging bias) and method 3	1.395	1.405
1964 to 1977	method 0 (without correction for aging bias) and method 4	1.327	1.285
1978 to 1984	method 0 (without correction for aging bias) and method 5	1.224	1.190
1985 to 1993	method 0 (without correction for aging bias) and method 6	1.021	1.018
1964 method change (20 percent more response)	method 3 and method 4	1.051	1.094
1978 method change (more complete response, nonresponse bias)	method 4 and method 5	1.084	1.088
1985 method change (vacancy imputation)	method 5 and method 6	1.195	1.160
One-month recall bias	method 5 and method 5 (without correction for recall bias)	1.103	1.096
Impact of vacancy imputation on vacancy nonresponse bias	method 0 (without correction for aging bias) and method 5 (without correction for recall bias)	1.110	1.086

TABLE 7 INDEXES OF TENANT RENT U.S. CPI-W AND NEW SERIES, 1940-2000, 1982-84 = 100		
December	BLS CPI-W, Rent of primary residence	New Series
1940	23.8	12.9
1941	24.7	13.5
1942	24.7	13.5
1943	24.8	13.7
1944	24.9	13.8
1945	24.9	13.8
1946	25.2	14.2
1947	26.9	15.7
1948	28.2	17.0
1949	29.3	18.1
1950	30.2	19.0
1951	31.7	20.6
1952	33.1	21.9
1953	35.0	23.8
1954	35.4	24.3
1955	35.9	24.9
1956	36.8	25.8
1957	37.4	26.5
1958	38.0	27.2
1959	38.6	27.9
1960	39.1	28.5
1961	39.6	29.2
1962	40.0	29.7
1963	40.4	30.2
1964	40.8	30.7
1965	41.2	31.2
1966	41.9	32.0
1967	42.8	33.0
1968	44.0	34.3
1969	45.6	36.1
1970	47.7	38.3
1971	49.5	40.4
1972	51.2	42.3
1973	53.7	45.1
1974	56.6	48.5
1975	59.5	51.9
1976	62.8	55.8

Table 7 continued		
Dec. to Dec.	BLS CPI-W	New Series
1977	66.9	60.7
1978	71.7	66.2
1979	77.4	72.8
1980	84.4	80.9
1981	91.5	89.4
1982	97.5	96.8
1983	102.2	102.7
1984	108.1	110.2
1985	115	117.8
1986	120.8	124.3
1987	125.3	129.5
1988	129.7	134.1
1989	135	139.7
1990	140.2	145.2
1991	144.8	150.0
1992	148.2	153.6
1993	151.6	157.2
1994	155.4	161.2
1995	159.3	165.2
1996	163.7	169.8
1997	168.8	175.1
1998	174.6	181.1
1999	179.9	186.6
2000	187.0	193.9

TABLE 8 OFFICIAL ESTIMATES OF CPI FOR RENT, REVISED CPI FOR RENT, AND GAP BETWEEN MEDIAN RENTS AND CPI, DECEMBER TO DECEMBER, 1930 TO 2001						
	1	2	3	4	5	6
	Original CPI-W for rent	Revised CPI	Change in median gross rent	Rental inflation gap, median vs. CPI	Revision	Revision as proportion of gap
1930-40	-2.7	-2.7	-2.4 *	0.4	0.4	1.00
1940-50	2.4	3.9	4.5	2.1	1.5	.68
1950-60	2.6	4.1	5.1	2.6	1.5	.58
1960-70	2.0	3.0	4.2	2.2	1.0	.44
1970-77	4.8	6.6	7.6	2.8	1.8	.62
1977-85:	6.8	8.2	8.5	1.8	1.4	.84
1985-95	3.3	3.4	3.6	0.4	0.1	.33
1995-2001	3.4	3.4	3.2	-0.3	0	.00
1940-1985	3.5	4.9	5.8	2.3	1.4	.61

Sources: Decennial Censuses of Housing, American Housing Survey, and CPI.

TABLE 9
COMPARISON OF CPI-U RENTAL INFLATION RATES WITH ALTERNATIVE RENTAL INFLATION MEASURES BASED ON AMERICAN HOUSING SURVEY MICRODATA, LOG PERCENT ANNUALIZED RATES

	Median gross rents, AHS	CPI-U, rent, IVQ to IVQ	Revised CPI-U, rent, IVQ to IVQ	Box-Cox Hedonic measure, AHS*	Repeat rent Measure, AHS*
1975-77	8.3	5.7	7.8	8.9	6.9
1977-79:	8.2	7.4	9.1	8.5	6.7
1979-81	10.9	8.3	10.2	10.7	8.6
1981-83	7.7	5.7	7.2	6.9	5.9
1983-85	7.2	5.9	6.8	7.0	not available
1985-87	4.6	4.4	4.7	5.4	4.2
1987-89	3.0	3.9	3.9	5.3	4.9
1989-91	4.3	3.5	3.6	5.7	5.0
1991-93	2.6	2.3	2.3	2.8	3.3
1993-95	3.6	2.5	2.5	3.9	3.6
1995-97	2.4	2.9	2.9	1.5	2.6
1997-99	2.7	3.1	3.1	4.7	3.6
1999-2001	4.4	4.2	4.2	3.2	4.2
Average Rate 1975-83	8.8	6.8	8.6	8.7	7.0
Average Rate 1985-2001	3.4	3.3	3.4	4.1	3.7
Average Rate 1975-2001	5.4	4.6	5.3	5.7	not available

Sources: American Housing Survey, CPI, and authors' calculations. CPI-U is CPI-W before 1978, when the CPI-U was introduced.
*The hedonic measure represents updated estimates based on the model in Crone et al, 2001. The repeat rent measure is estimated from the AHS panel using the standard repeat sales/rent methodology. Since the AHS panel changed in 1985, the repeat rent measure could not be estimated for the 1983 to 1985 interval.

TABLE 10
COMPARISON OF ALTERNATIVE RENT PRICE INDEXES WITH OTHER PRICE INDEXES, LOG PERCENT ANNUALIZED INFLATION RATES
(Underlying data are annual average price levels.)

		1940 to 1985	1985 to 2001	Difference
Official rent estimates	CPI-W, not seasonally adjusted, tenant rents, BLS	3.43	3.37	-0.06
	PCE chained price index, housing services: tenants, BEA	3.62	3.45	-0.17
	PCE, chained price index housing services: owners equivalent, BEA	3.59	3.52	-0.07
New rent estimate	Adjusted CPI-W rents, new estimates	4.84	3.46	-1.38
Median rents	Median gross rents, Census and American Housing Survey, Census Bureau	5.78	3.45	-2.33
Residential structures	Residential fixed investment chain price index, BEA	5.06	3.15	-1.91
Aggregate price measures	CPI-W all items excluding shelter, BLS	4.50	2.81	-1.69
	PCE chained price index, BEA	4.39	2.64	-1.75
	GDP chained price index, BEA	4.37	2.40	-1.97
	PPI all items, BLS	4.51	1.64	-2.87
Wage measure	Average Hourly Earnings, manufacturing, BLS	6.40	2.82	-3.58

Sources: U.S. Bureau of Economic Analysis, U.S. Bureau of Labor Statistics

TABLE 11
COMPARISON OF REAL HOUSING SERVICES ESTIMATES WITH ALTERNATIVE REAL GROWTH MEASURES
(year average data)

		1940 to 1985	1985 to 2001	Difference
Housing services	Real PCE housing services, BEA	4.63	2.44	-2.19
	Real PCE housing services adjusted, new estimates	3.52	2.49	-1.04
Residential net stocks	Real net stock of residential fixed assets, BEA	2.93	2.54	-0.39
Residential investment	Real residential fixed investment, BEA	3.78	2.56	-1.22
Aggregate activity	Real GDP, BEA	3.93	3.05	-0.87
	Real PCE, BEA	3.71	3.31	-0.39
Demographic	Nonfarm Payrolls, BLS	2.45	1.88	-0.56
	Population, Census Bureau	1.31	1.13	-0.18
	Households, Census Bureau	2.02	1.38	-0.64

Sources: U.S. Bureau of Economic Analysis, U.S. Bureau of Census

Figure 1. Comparison of Indexes based on Six-Month and One-Month Rental Increases