WORKING PAPERS
RESEARCH DEPARTMENT

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MEASURING AMERICAN RENTS:
A REVISIONIST HISTORY

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July 2001

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ABSTRACT

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Until the end of 1977, the method used to measure changes in rent of primary residence in the U.S. consumer price index (CPI) tended to omit price changes when units changed tenants or were temporarily vacant. Since such units typically had more rapid increases in rents than average units, omitting them biased inflation estimates downward. Beginning in 1978, the Bureau of Labor Statistics (BLS) implemented a series of methodological changes that reduced this bias. We use data from the American Housing Survey to check the success of the corrections. We compare estimates of the historical series adjusted for the BLS changes in methodology with a new hedonic estimate of changes in rental rates. We conclude that from 1940 to 1977 the CPI for rent would have been about 60 percent higher if current BLS practices had been used -- between 1.3 and 3.5 percentage points. Even after the corrections have been made, our hedonic estimates suggest that the current CPI methodology may still understate the rental inflation rate by one-half to 1 percentage point.

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MEASURING AMERICAN RENTS:
A REVISIONIST HISTORY

I. Introduction and overview

Before 1978 the data used to estimate rental inflation for the U.S. consumer price index (CPI) suffered from two forms of downward bias: aging bias and nonresponse bias. Aging bias occurs because the quality of the average rental unit tends to deteriorate over time because of inadequate maintenance. If the rental price of a unit remains constant and its quality deteriorates, its quality-adjusted rent has risen. Therefore, rental inflation data unadjusted for aging bias is downwardly biased.

Nonresponse bias, the more important of the two biases and the focus of this paper, has two sources: (1) apartments become vacant and hence there is no rent information available and (2) apartments change tenants and BLS price inspectors lose contact with tenants, preventing collection of rental data. Since changes in tenancy normally coincide with rental price increases, ignoring nonrespondents may result in a large downward bias. Only the vacancy part of nonresponse bias has been explicitly studied by the Bureau of Labor Statistics (BLS), and the impact of vacancy nonresponse bias and the imputation to correct this bias has not been discussed in either Moulton's review of rental inflation or Stewart and Reed's current methodology research series.[1]

Repeated investigations have suggested that prior to 1978, the CPI rental index was downwardly biased. (Ozanne, Humes and Schiro, and Lamale). Between 1940 and 1977, a

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[1]Stewart and Reed suggest that the only adjustment needed to pre-1978 data is an adjustment for aging bias. We believe that an adjustment is needed for the nonresponse bias as well.

period during which the methodology underlying the index was most vulnerable to nonresponse bias and was uncorrected for aging bias, the CPI for rent rose 2.8 percent annually (Table 1). Bureau of Census measures of rent, reported in the decennial Census of Housing and the biennial American Housing Survey, show that median gross rent rose 5.5 percent annually -- 2.7 percentage points faster than the CPI for rent. If we take the CPI data at face value, this implies that the quality of the median rental unit increased 2.7 percent a year during this period. By comparison, from 1930 to 1940 and from 1983 to 1997, median gross rents rose less than half a percentage point faster than the CPI rent index, implying a substantially lower increase in quality.[2] This anomaly is explained in part by the downward bias in the CPI rental increase due to nonresponses.

Section II of this paper discusses the nature of nonresponse bias in the rental CPI and attempts by the BLS to correct it. Section III presents our estimates of rental inflation using hedonic techniques and compares these hedonic estimates with estimates adjusted for changes in the BLS methodology. Section IV summarizes the major conclusions of the paper.

II. The Nature of Nonresponse Bias and Attempts to Correct It

All sample surveys suffer from nonresponse, i.e., incomplete returns from some part of the targeted sample. This was not a major problem in the BLS rental survey prior to 1942 when price inspectors obtained their data from the files of real estate agents and large-property owners. This system had the advantage of avoiding a relationship with the tenant. The price inspector could directly compare current rents with past rents, regardless of whether the tenant had

[2] Prior to 1940, the BLS directly interviewed landlords rather than tenants, and it believes the problem of nonresponse bias was not a major one.

changed. If a unit was vacant, a comparable unit could usually be found from the books.

Changes in BLS methodology starting in 1942 introduced serious nonresponse bias into the rental-price series. Price inspectors were instructed to obtain the rents from the tenants directly rather than from the records of landlords or real estate managers.[3] Roughly 30,000 tenants were sampled. This typically involved an initial interview to elicit cooperation and gather data about the unit. After the initial interview, the tenant was mailed a questionnaire *quarterly*. The price inspector would report rental increases (called price relatives), and the recorded rate of rental inflation would reflect the average rate of rental increase. Approximately 50 percent of the initial mail questionnaires were returned completed by the tenant, and an additional 20 percent were returned upon follow-up. But 30 percent of the mail questionnaires were not completed, a relatively high rate of nonresponse.

Rents are usually increased annually, and such increases are typically associated with lease renewals, a time when tenants are most likely to move. When the tenant moves, the contact between price inspector and tenant is broken, and the rental increase goes unrecorded. This link between rental increases and change of tenants or vacancies biases downward the average rate of rental increase -- the rent quotations missed are precisely those that show increases.

Between 1952 and 1994, the BLS largely corrected the biases in the CPI in five steps. However, to our knowledge, the extent of this problem has never been investigated. We estimate the effect on the bias of these changes by the BLS and adjust the historical rental inflation for the

[3]An important impetus for this change was the implementation of wartime rent controls. It was feared that rental increases that evaded or violated rent control laws might not be accurately reported by real estate agents or landlords. By gathering data on the terms of the rental agreement, the price inspector would be able to detect changes in the terms, such as requiring the tenant to pay for utilities that had previously been included in the rent.

change in methodology. The five steps included:

(1) a reduction in the frequency of collection of prices from quarterly to semiannually in 1952;

(2) a major change in sampling procedures and methodology in 1978 that resulted in a significant reduction of the number of nonrespondents but introduced a recall bias in the estimate;

(3) an adjustment to the rental component of the CPI in 1983 that attempted to correct for vacancy-related nonresponse bias;

(4) an aging-bias adjustment based on Randolph's (1988a and b) methodology;

(5) elimination in 1994 of the recall bias that was introduced in the 1978 changes.

We will discuss each of these changes in order and estimate their effects.

II.1 The reduction in the frequency of rental data collection (1952)

The importance of the frequency of rental data collection for the size of any nonresponse bias is based on certain characteristics of the U.S. rental market. First, changes in rents are periodic; rent typically increases yearly, often at the time the lease is renewed. Data from the Property Owners and Managers Survey for an anonymous city in 1993 showed that 43.5 percent of all units had annual leases, 2.3 percent had leases longer than one year, 39.3 percent had leases less than one year, and 14.6 percent had no leases (Genesove). Second, a large proportion of rental units change occupants every year, and because a unit's rental price is not controlled by a lease when the tenant changes, its rent is likely to rise. Genesove reports 34.9 percent of all U.S. rental units in the American Housing Survey from 1975 to 1981 turned over each year. When tenants are the source of rental data, the relationship between tenant and the price collector is typically broken when the tenants change and must be renewed, often causing the rental data to be omitted.[4] Third, rental increases tend to be lower for tenants that continue than for new tenants. This is probably the result of two effects: an unexpectedly high rental increase is more likely to induce a renter to move, and the existing rental price may be a focal point in bargaining between the tenant and the landlord (see Genesove for a discussion).

Given these processes for rental adjustment, we can derive a general formula for the size of bias from a periodic survey of rents. We assume that prices are collected n times a year from each tenant in the sample, and that on average the price is increased once a year. The probability that a price increase occurs in any given sampling is 1/n.

To illustrate, let us suppose that rental units are continuously occupied and the relationship between the price inspector and the tenant is never broken, and that the annual rent increase is π_c. The price inspector then will record a zero increase n-1 times, and a rental increase of π_c one time. In annualized terms, the rate of rental increase is zero n-1 times and $n\pi_c$ once, for an average annual rate of growth of $n\pi_c/n = \pi_c$. To be concrete, let n be 4 and π_c be 4 percent. Then in a given year, 3 increases of 0 percent are recorded, and one of 4 percent. Annualizing the rates, we have three quarterly readings of 0 percent and one of 16 percent, so that the average annualized rate of increase is 16 percent divided by four, or 4 percent.

Now we introduce two complications to this simple scenario: some tenants leave at the end of their annual lease period, when the rent is increased, and the units from which tenants depart have, on average, a higher rate of increase than the units of continuing tenants.

[4]Moreover, when the tenant changes, a vacancy occurs, typically lasting one or two months, which also contributes to a break in the rent collection series.

On the date the rent rises, the tenant leaves with probability t and no price increase is observed because the relationship between the tenant and the price inspector is broken, and with probability 1-t the tenant continues and the price increase is observed. If the tenant continues, the rent increase for that period, at an annualized rate, is $n\pi_C$. If the tenant does not continue, the rent increase is $n(1+a)\pi_C$, but this price is not observed since the tenant exits the sample.[5] The true rate of inflation is $\pi = (1+ta)\pi_C$. Price inspectors record n-1 observations of zero, 1- t observations of π_c, and do not obtain an observation t times. The total number of observations is n-t, and the sum of the annualized price increases recorded is $(1-t)n\pi_c$. So the average recorded price increase is $\pi_C(1- t)/(1 - t/n)$.[6] Clearly, this is biased downward from $\pi = (1+ta)\pi_c$.

The turnover rate t varies but has been about one-third overall. That figure does not include vacancies, which have averaged about 7 to 8 percent. So if we include vacancies, turnover (t) is about 0.4. According to data in Rivers and Sommers (1983), the average increase for new tenants is about 1.3 times the increase for continuing tenants, so a = 0.3. From 1942 to 1952, data collection was quarterly (n=4), so the observed rate, according to this model, would be $.595\pi$, with a nonresponse bias of $.405\pi$. From 1953 to 1977, data collection was semiannual (n=2), and the observed rate of rent increase would have been $.670\pi$ and the theoretical nonresponse bias would have been $.330\pi$. Thus the change from quarterly to semiannual

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[5](1+ a) is the ratio of the rental increase for a new tenant relative to the increase for a continuing tenant.

[6]There is an additional factor that complicates the analysis. The hazard rate of tenant turnover decreases over time: a tenant who has been in residence for k years is more likely to renew than one who has been in residence less than k years. Among other effects, this can impart a dynamic survivorship bias because a fresh sample will behave differently from an aged sample. Thus the change of methodology in 1978, discussed below, as well as that in 1942, may have influenced the measured inflation rate.

collection of rental data reduced nonresponse bias by .075π.[7]

The method of survey by mail was eventually deemed unsatisfactory because of the large number of nonrespondents, and in the 1964 revision to the CPI, the BLS instituted a system of using part-time agents to collect rental data by personal visit or telephone. Forty thousand units were surveyed semiannually to obtain a total of 80,000 prices annually, or an average of 6,667 per month. Rental units were still priced every six months. No substitution was permitted for units whose prices were not obtained. For a short period in 1964 the data were collected using both the old and the new survey methods for comparison purposes. During this period, there was very little difference between the two series. By the end of the overlap period, June 1964, the revised index for rent was 107.8 (on a basis of 1957-59 = 100) compared with the unrevised index of 107.9, so the revised index rose more slowly. The June 1963 rent index was 106.8, so the rental CPI at this time was rising at an annual rate of about 1 percent. Thus it does not appear that the 1964 revision did much to eliminate nonresponse bias.

II.2 Major changes in estimating rental inflation in 1978

Beginning in 1978, a new survey method was instituted. The number of rental units surveyed was reduced substantially to 18,000. The intention was to ensure that the sampling of rental units was as thorough as possible and, in particular, to capture rent increases when the tenant moved. Data were also obtained on the length of occupancy of new tenants. Price inspectors could choose to interview the landlord or manager instead of the tenant and typically

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[7]It is difficult to be entirely sure when changes in procedure took place during the 1950s and 1960s, because BLS documentation was less complete during this period. The BLS (1966) suggests that the 1954 revision had changed rental price collections to twice a year. Moreover, an example in the discussion of the 1953 revision to the CPI also suggests that collections had been changed to every six months.

did so. Price inspectors were to reinterview the tenant, manager, or owner of the unit every six months. Nonresponse fell to less than 14 percent.

In addition, a new method was instituted for using the rental data obtained from the interview. First, respondents were asked the level of last month's rent as well as the current month's rent. Then two comparisons were made: the six-month price increase using the previous interview and the one-month price increase. The rental index was computed using both the one-month change and the six-month change, weighted so as to minimize fluctuations. Defining I(t) as the level of the index at month t, and Rt,t-k as the change in rent from k months ago, the rental formula was:

$$I(t) = .65\ R_{t,t-1}\ I(t-1) + .35\ R_{t,t-6}\ I(t-6). \qquad (1)$$

A study of the post-1977 data by two BLS economists, Joseph Rivers and John Sommers, revealed that the BLS rental price estimates still suffered from two biases: recall bias and the vacancy component of the nonresponse bias. Recall bias was a systematic tendency for one-month price changes to be less than the sixth root of six-month changes; six-month changes had the advantage of being based on previous records and not the recall of the tenant or landlord. The 1978 changes eliminated most of the component of nonresponse bias associated with new tenants.

Evaluation of 1978 CPI revisions using revision overlaps. The empirical evidence on the effect of the 1978 changes is stark. From January to June 1978 the BLS conducted the rental survey using both the original and the revised methods. The main purpose of the overlap period was to allow for the calculation of wage rates that were indexed to the old series, but the overlap gives us a window onto the change due to the revision that we can compare to the theoretical model. In the overlap period the rent index using the pre-1978 methodology shows a six-month increase of 3.2 percent while the index using the post-1978 methodology shows a six-month increase of 3.5 percent. This reduction of 0.3 percentage point is roughly 10 percent, so we have a reduction of $\pi/10$ or about half the bias we estimated from our model given the frequency of sampling and the average turnover rate. Since the new methodology was implemented in stages over three months, this may represent an understatement of the adjustment.[8] We believe that reducing the nonresponse bias adjusted the reported inflation rate upward by 24 percent, while the downward recall bias took back 9 percent, for a net change of 15 percent.

II.3 Adjustments to correct for the vacancy-related nonresponse bias (1983)

Vacancies present a special problem in collecting rental data because a unit that is vacant at the time of a scheduled interview will not have a recorded rent to compare with the previous time or the next time it is collected. Therefore, no increase can be computed for the unit over that period. Although the 1978 procedures had reduced nonresponses from 30 percent to 13.6 percent, nonresponses due to vacancies were little changed and now accounted for half of all nonresponses. If rental increases for units that become vacant are higher than the average rental increases, there is a negative nonresponse bias associated with vacancy.

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[8]The new CPI procedure was introduced in a rolling fashion. Different cities had their rents recorded in different months of the quarter, and some did not begin reporting data until February of 1978. Thus by the termination of the six-month overlap period, some cities had reported under the new procedure for only four months. Indeed, all of the deviation between the old and the new data occurs between April and June. These numbers are for the CPI-W, revised and unrevised. Seasonally unadjusted rent levels for the unrevised CPI (W) were, from December, 1977 to June, 1978: 157.9, 158.7, 159.7, 160.6, 161.4, 162.2 and 163.0. For the revised CPI-W, they were, from January 1978 to June 1978, 158.8, 159.7, 160.5, 161.4, 162.6, 163.5. For the revised CPI-U they were, for the same period, 158.8, 159.7, 160.5, 161.5, 162.7 and 163.6. These data were published in the CPI monthly detailed reports for those months, and then reviewed by Layng.

Using CPI rental microdata to estimate nonresponse bias resulting from vacancies. Rivers and Sommers use all the CPI rental price microdata observations collected by price inspectors from April 1979 to March 1981 to get a measure of the bias associated with vacancies.[9] In this period, there were 56,510 interview attempts, from which 48,809 good interviews resulted (86.4 percent). Reasons for noninterviews were vacancies (3,833, or 6.8 percent), no one at home (2,619, or 4.6 percent), refusal (745, or 1.3 percent) and other (504, or 0.9 percent). However, only 45,758 six-month changes were recorded for the 48,809 units with good interviews. Presumably a good interview was conducted 3,051 times at a rental, but no six-month price change was recorded because six months previously no price data had been obtained for that particular unit.

Rivers and Sommers divided their good interview sample into continuing tenants (those with six or more months of occupancy, 81.2 percent of the sample) and new tenants (18.8 percent). This breakdown is consistent with a turnover rate of about 35 percent annually and, therefore, suggests that the new survey did succeed in capturing new tenants. As reported in Table 2, the annual rate of increase in rents for new tenants was 20.4 percent. In contrast, 46 percent of *continuing* tenants experienced an increase that averaged 8.5 percent annually. The average rental price for all units increased at an annual rate of 10.7 percent. This is consistent with the view that, on average, rental prices are raised once a year and that rental increases are greater for new tenants. If the survey had captured only continuing tenants, the average rate of rental increase would be underestimated by 2.2 percent, or just larger than the 1/5 that theory

[9] In addition, they used additional microdata back to January 1979 for the purpose of calculating six-month changes.

suggests. After 1977, it appears that the nonresponse bias associated with tenancy change had been eliminated and that the 1978 revision had effectively reduced the nonresponse bias to vacancy bias.

During the period from April 1979 to March 1981, the vacancy rate for surveyed units was 6.8 percent. If we substitute the vacancy rate for the turnover rate in equation (1), we obtain a theoretical vacancy bias of roughly $\pi/30$. Finally, if vacancies have the same high rate of rental price increase as apartments with new tenants, then the true rate of rental inflation between April 1979 and March 1981 would have been 11.3 percent annually rather than 10.7 percent. The bias induced by vacancy omissions by this measure is 0.6 percent, or roughly .05 π.

By separating responses into those of new tenants (less than six months' occupancy) and continuing tenants, Rivers and Sommers showed that new tenants had higher rates of price increase than continuing tenants. As shown in Table 2, 46.4 percent of continuing tenants experienced rent changes in the previous six months, while 80.6 percent of new tenants experienced rent changes. Moreover, those new tenants who experienced rent changes experienced higher rates of rent increase (12.1 percent) than continuing tenants (8.9 percent). Using this information, the BLS developed a correction for the vacancy bias in 1983, which involved the estimation and imputation of expected rents for vacant units. This change in methodology probably accounted for another 9 percent upward adjustment to rental inflation, resulting in a total nonresponse adjustment of 0.33 times the rental inflation rate.

II.4 The adjustment for aging bias (1988)

None of the changes to the BLS method in 1978 and 1983 to correct for nonresponse and vacancy bias addressed the issue of aging bias in the estimate of rental inflation. BLS economists have long worried about aging bias, but it was not until the late 1980s that they were satisfied that they could estimate it accurately. Aging bias refers to the underestimation of rental increases because of the systematic deterioration in the quality of housing services provided by a rental unit as it ages. Historically, the BLS has adjusted the change in rent for *observed* quality changes, such as the addition of a room. But prior to 1988 the agency did not correct for the systematic deterioration in quality associated with aging. If a unit deteriorates systematically with age, a constant rent over the six-month period implies an increase in rent on a quality-adjusted basis.

There are two potential problems in estimating the effect of physical deterioration on rents. The first is the so-called vintage effect. This effect arises when there are quality characteristics other than physical deterioration associated with age but not other measured characteristics of the residence. For example, the more extensive use of insulation in houses built after the 1970s would raise the unmeasured quality of those units. On the other hand, units built prior to World War II and still occupied may represent the highest quality units built in those years based on the assumption that the lower quality units built at that time are no longer in use. These so-called vintage effects make it difficult to get an accurate estimate of the effect of physical deterioration on rent. The second problem in estimating the effect of aging on rent is that units of different types (e.g., apartments versus detached houses) may deteriorate at different rates. In his 1988 article William Randolph (1988b) was satisfied that he had solved both of these problems in estimating the effect of systematic physical deterioration on rents.

Randolph argues that including a sufficient number of housing and neighborhood characteristics in a hedonic equation would render the remaining vintage effect minimal. He included housing characteristics like the presence of a dishwasher or washer/dryer and neighborhood characteristics like the percent of the population with a college education. He also estimated different aging effects depending on the number of rooms in the unit, whether the unit was detached, and whether it was rent controlled. His resulting estimate of the *average* effect of aging on rent was -.36 percentage point a year and did not vary with the inflation rate. The BLS has used this estimate of the effect of aging to adjust the rent component of the CPI since 1988. This adjustment increased the rental inflation rate by 9 percent.

II.5 Elimination of the recall bias (1994)

The recall bias problem introduced in 1978 was solved in 1994 when the BLS discontinued the use of reported one-month rent increases in estimating rental inflation (Armknecht, et al.).

The data reported by Rivers and Sommers illustrate the recall bias. Overall, 24,182 six-month changes were reported between April 1979 and March 1981, but only 2,541 one-month changes. The number of reported one-month changes is just 63 percent of the 4,030 expected based on the number of six-month changes. This suggests that a large percentage of one-month changes are not being recalled or reported.

The average one-month change for all tenants cannot be fully derived from the data in Rivers and Sommers, because one-month changes for tenants with less than six months' occupancy were not given in detail. We estimated the one-month rent changes for those tenants by establishing an upper and lower bound and taking an average. We assume that the lower bound for new tenants was the average one-month rent change for tenants with six months' or more occupancy (10.22 percent). This assumption is based on the fact that new tenants consistently had higher six-month rent increases than tenants with six months or more occupancy. It is reasonable to assume, then, that the one-month change for new tenants was at least as high as the one-month increase for long-term tenants. The upper bound for one-month changes for new tenants (less than six months) is the highest six-month change for any occupancy group. According to the data in Rivers and Sommers, those with one-month occupancy had the highest six-month change (13.29 percent). The average of the upper and lower bounds for one-month changes for new tenants is 11.76 percent (Table 3).

The average *annual* rent change implied by the one-month changes in 1979-81 was 7.5 percent compared to a 10.7 percent rate of increase in the six-month changes. Thus the recall bias of the one-month change compared to the recorded six-month change was .29 π. However, the impact of the recall bias on the measured inflation rate is less than this, since the rental index was computed using both the one-month rate and the six-month rate.

What is the quantitative impact of a given recall bias on measured rental inflation? Suppose the true *monthly* inflation rate is π. The six-month rental inflation rate will be $(1+\pi)^6$. If the one-month recall bias is e, then the reported one-month change will be $(\pi - e)$. The formula given in equation (1) to compute the rental index can then be written as the following sixth order difference equation:

$$I(t) = .65(1+\pi-e)\, I(t-1) + .35\, (1+\pi)^6\, I(t-6). \qquad (2)$$

If we assume that measured monthly inflation in the steady state equals

$1 + \pi - de$

where

d = the impact on the measured inflation rate of the recall bias e.

Then

$$I(t) = (1+\pi\text{-}de)\, I(t\text{-}1) \quad \text{and}$$

$$I(t) = (1+\pi\text{-}de)^t\, I(0). \qquad (3)$$

To compute d we substitute and obtain:

$$(1+\pi\text{-}de)^t\, I(0) = .65(1+\pi\text{-}e)(1+\pi\text{-}de)^{t-1}\, I(0) + .35\,(1+\pi)^6\,(1+\pi\text{-}de)^{t-6}\, I(0) \qquad (4)$$

Dividing through by $(1+\pi\text{-}de)^{t-6}\, I(0)$ and subtracting the right-hand side, we obtain:

$$1 - .65(1+\pi\text{-}e)/(1+\pi\text{-}de) - .35\,[(1+\pi)/(1+\pi\text{-}de)]^6 = 0 \qquad (5)$$

Now, performing the division indicated by the second term on the left-hand side of equation (5):

$$(1+\pi\text{-}e)/(1+\pi\text{-}de) = 1\text{-}\ e\,(1\text{-}d) + \text{error}\ . \qquad (6)$$

The remainder from the division is actually $((\text{-}e\,(1\text{-}\ d))/(1+\pi\ \text{-}\ de)$. But both π and e are assumed to be much smaller than one and d is less than one. Therefore, the remainder can be approximated by -e (1-d) plus a small error, where the error is on the order of π times e.

Performing the division indicated by the third term on the left-hand side of equation (5):

$$(1+\pi)/(1+\pi\text{-}de) = 1 + de + \text{error} \qquad (7)$$

The remainder from the division is actually $de/(1 + \pi - de)$, but for the reasons mentioned above, this denominator is very close to one, and the remainder can be expressed as de plus a small error, where the error is on the order of π times e. Ignoring the error and raising the right-hand side of equation (7) to the sixth power, we obtain

$$(1+de)^6 = 1 + 6de + \text{error} \qquad (8)$$

where the error represents all the exponentiated values of de and is therefore very small.

Ignoring the error terms and substituting the right-hand sides of (6) and (8) into (5), we have approximately

$$1 - .65\,(1 - e(1\text{-}d)) - .35\,(1 + 6\,de) = 0$$

or

$$d = .2364. \tag{9}$$

This implies that if the one-month recall bias is $.2\pi$, the measured inflation bias will be $.047\pi$.[10]

In the period from 1978 to 1981, the measured rental inflation, by these calculations, should have been 10.1 percent -- lower than the 10.7 percent six-month rate by about one-fourth of the 2.9 percentage point recall bias. In fact, during this period the CPI for rents rose only 9.1 percent, which is lower than the Rivers and Sommers data suggest it should have been.

When recall bias was corrected in 1994, the impact on the rental index was estimated at 0.22 percentage point, or about $.09\ \pi$ (at the time the rental inflation rate was about 2.5 percent).[11] For the impact of recall bias to be this large, given that d is .24, the one-month rate should have been 40 percent lower than the six-month rate. The Rivers and Sommers data suggest that the one-month estimate was 29 percent less than the six-month rate, and therefore, recall bias should have been only $.07\ \pi$. Thus, it seems possible that the recall bias has changed somewhat over time. Since the inflation rate fell considerably from the period of the Rivers and Sommers data (1979-81) to 1994, some impact on the recall rate would not be surprising.

II.6 Total impact of BLS adjustments

Table 4 presents our estimates of the impacts of the BLS methodological changes on

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[10]A simulation over a six-ear period with a = .005 and e = .001, so that the annual inflation rate is about 6 percent, yields d = .2362.

[11]In 1994 the BLS abandoned the use of a weighted average of six-month and one-month changes to estimate rental increase. Since then the Bureau has used the sixth root of the six-month change to estimate the one-month change.

rental inflation rates. The estimates of the impact of increased response rates for new renters and of vacancy imputation are our estimates, while estimates of aging bias and recall bias are from the BLS. In 1999 Stewart and Reed published an adjusted CPI that incorporated the adjustments for 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. The total impact of the corrections on the rental inflation rate was roughly 0.4 times the rental inflation rate plus 0.36 percentage point, from 1942 to 1952, and .33 times the rental inflation plus 0.36 percentage point, from 1952 to 1977. Prior to these corrections, 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 that ranged between 1.3 and 3.2 percentage points a year.

To evaluate the adequacy of our adjustments to the rental CPI, we used hedonic regression techniques and data from the American Housing Survey to create an independent index of rental housing from 1975 to 1995. Our hedonic estimates based on the American Housing Survey suggest that there may still be some downward bias in inflation rate for rent as reported in the CPI.

III. Measuring Rental Inflation Using Hedonic Estimation Techniques

Housing is essentially a bundle of goods: kitchen, bathrooms, bedrooms, etc. There is a vast literature on hedonic techniques applied to the housing market to estimate the underlying prices of various elements of the housing bundle (see Sheppard for a review and references therein for reviews of the empirical literature). There is almost as large a literature devoted to constructing indices of house price appreciation, and many of these papers use hedonic techniques to control for changes in house quality over time (see Malpezzi, Chun, and Green for a recent example). Other than Thibodeau (1995), only a few papers measure rental price increases using hedonic techniques, and these have tended to focus on metropolitan rents, rather than the national rate of inflation (see references in Thibodeau, 1992 and 1995.) In this paper, we use 11 cross-sections of the American Housing Survey spanning 1975-95 to construct a price index for rental housing that provides a basis for evaluating the longer term accuracy of the CPI for rental housing as well for analyzing the impacts of adjustments to the series over the sample period.

A constant-quality rental price index constructed using a hedonic regressions differs from the consumer price index in practice, but not in principle. In practice, the current CPI for rents holds quality constant by (1) correcting for aging bias, (2) either omitting units whose characteristics have changed (for example, by the addition of air conditioning) or, where available, pricing out the changes in characteristics, and (3) using relative rent increases only from unchanged tenant unit locations. Our hedonic regressions, on the other hand, systematically price out all available differences in characteristics, including location.[12] Thus, in addition to characteristics of structures, units, and rental terms, our hedonic analysis also includes neighborhood and geographical characteristics (region, urban-rural status, and central city location) to control for location.

To construct measures of the rental inflation rate, we estimate the market rental prices of

[12]In principle, some neighborhood characteristics can change over time, and as a result, the quality of housing at an unchanged location may change. However, in practice such changes in neighborhood characteristics are too small and infrequent to have a significant impact on the overall rate of inflation.

the component housing traits, and using the estimates of the stock of these traits, we can estimate the change in the rent of an average constant quality rental unit. We specify the dependent variable in our estimation as a Box-Cox transformation of rent so that the hedonic regression takes the form:[13]

$$\frac{R_{it}^{\lambda_t} - 1}{\lambda_t} = \beta_t X_{it} + u_{it} \tag{10}$$

where:

R_{it} is the rental rate of unit j in time t;

X_i is a k element row vector of housing traits of house i of I houses;

β_t is a vector parameters associated with individual traits; and

λ_t is the Box-Cox transformation parameter.

If b_t is our estimate of β_t, then $(\lambda_t b_t X_{it} + 1)^{1/\lambda}$ is an estimate of rent for house i at time t.

Using estimates of the parameters of (10), we can construct indexes of monthly rents as follows: Let $W_{it} = Z_{it}^{-1}$ where Z_{it} is the sampling probability of house i. Also, let X_{it} be an I by k matrix whose rows consists of values of each of the housing traits for the i^{th} house of the I rental units in the sample; and W_{it} be a one by I vector of weights that blows the sample up to the universe. Then $W_{it}\,(\lambda_t b_t X_{it} + 1)^{1/\lambda_t}$ is a measure of the nominal value of rental services in period t in dollars of period t. The change in the nominal value of housing services from t to t+n is given by $W_{it+1}\,(\lambda_{t+1} b_{t+1} X_{it+1} + 1)^{(1/\lambda_{t+1})} / W_{it}\,(\lambda_t b_t X_{it} + 1)^{(1/\lambda_t)}$. Holding the matrix of characteristics of homes

[13]There is a large literature on the appropriate choice of functional form for the hedonic price function (see Linneman 1980, for example). The Box-Cox transformation nests both linear ($\lambda = 1$) and semi-log ($\lambda = 0$) models.

constant, we can determine the price of the same bundle of services in period t+n by $W_{it}\,(\lambda_{t+1}b_{t+1}X_{it} + 1)^{(1/\lambda_{t+1})}$. We will construct a Laspeyres price index of rental services, $L = W_{it}\,(\lambda_{t+1}b_{t+1}X_{it} + 1)^{(1/\lambda_{t+1})}/W_{it}\,(\lambda_t b_t X_{it} + 1)^{(1/\lambda_t)}$, using biennial data, so that n=2. Similarly, we can construct an analogously defined Paasche price index, P, and a Fisher ideal index, $F = (L*P)^{0.5}$. In the results that follow we focus on the Fisher index that is chained together across years.

Data. The American Housing Survey national cross-sections are useful for evaluating changes in the price of U.S. rents for two reasons. First, they have data on housing attributes, and rental rates that can be used to estimate hedonic equations. Second, each cross-sectional sample has associated weights that can be used to expand the sample to the housing universe. These weights allow the calculation of the total flow of rents, given a set of estimated trait prices.

There are, however, a number of problems with the AHS data, one of which is missing values. Although every observation in the AHS sample has an associated weight that can be used to expand the sample to national totals, some observations have missing values for the key variable, such as rent, for which we wish to impute national totals. For those observations with missing rents, we impute the rental value using the estimated rental equation. A small number of observations in a few cross sections have missing data on housing or neighborhood traits. In these cases, we set the value of the trait to zero and include a dummy variable in the regression, indicating a missing value to capture any systematic differences in houses associated with missing values on the trait.[14] Truncation presents another problem in the AHS data. The rent

[14]In most regressions, the coefficients on these dummy variables are highly insignificant, and the variables were excluded from the final regressions.

data have upper bounds on their values, and these upper bounds change across years. Matching truncation limits across years has virtually no effect on our hedonic-based indexes, and the reported results do not include any corrections for truncation.

Another problem with the American Housing Survey is that there are really two separate panel data sets involved. Data from 1975-83 is based on the first panel while the data from 1985-95 are from a new panel. Not only are the samples different in the two periods, but the survey questions differ across samples as well. Moreover, in the earlier period, there were differences in the survey from year to year. These changes limited the number of variables that could be used in any pair of years. In the latter period, there was very little change in the survey from year to year. There are two main consequences of these changes in the AHS survey. First, models change from one pair of cross-sections to the next in the first part of the sample. We do not think this has any appreciable effect on the hedonic estimates. Second, our estimates of inflation for the two years 1983 to 1985 are suspect. The set of regressions spanning the two samples, 1983-85, gives what is probably the least reliable estimate of rental inflation because of changes in the sample and the survey questions.

Table 5 displays the sample means and standard deviations of the variables used in the analysis for the 1975 and 1985 and 1995 cross-sections.[15] As is evident from Table 5, our data include a rich set of structural, unit, and neighborhood characteristics as well as information on rental terms and geographic location. Examination of the unit characteristics indicates that the quality of units is improving over time: the number of rooms, bathrooms, the presence of central air conditioning, and satisfaction with the unit are all increasing. Negative measures of quality --

[15]Means of the dummy variables for missing values are available on request.

holes in the floors and presence of mice -- are decreasing. Neighborhood characteristics, on the other hand, appear to worsen slightly over time: concern with crime and noise increases and satisfaction with the neighborhood decreases slightly. One particularly noteworthy fact is that mean building age rises substantially over the course of the 20 years, from nearly 27 years to more than 38 years, so that sampled rental units are increasingly in older buildings and probably in older communities.

Hedonic estimates based on equation 10 are estimated for the 11 biennial cross-sections from 1975 to 1995. Table 6 presents results for the 1975, 1985, and 1995 cross-sections. The estimated coefficients (trait prices) are generally of the expected signs and of reasonable magnitudes. The relative prices of individual traits are generally consistent across time periods. Note that the adjusted R square declines over time, indicating greater variance. In particular, the depreciation variables, age and age squared, become quantitatively less important.

Changes in the Box-Cox transformation parameter, λ, over time warrants additional discussion. Table 7 presents these parameter estimates for the 11 cross-sections.[16] The λs increase over time, from 0.38 in 1975 to 0.61 in 1995. The hypotheses that $\lambda = 1$ or $\lambda = 0$ can always be rejected, indicating that neither the commonly used semi-log hedonic specification or the linear specification is appropriate. The increase in λ over time indicates that the distribution of rents is becoming less skewed over time as is clear in Figures 1a and 1b; the top graph in each figure shows a histogram of actual rent for either 1975or 1995. The second pair of graphs

[16]Because the variable set changes slightly through time, two equations were estimated in some years, reflecting the traits data available for the previous or following cross-sections. The estimates of λ are virtually identical in all cases where two estimations were made on a cross-section.

corresponds to predicted rents from semi-log estimations and from the Box-Cox model. Note that the semi-log predictions are substantially more skewed than actual rent in 1995.

Note that changes in λ_t across time periods introduces a bias into the price change index because changes in λ_t change the measure of central tendency. Our estimates of rent, $\Sigma_i(\lambda_t b_t X_{it} + 1)^{1/\lambda_t}/I$, are an unbiased estimate of mean rent only if $\lambda_t = 1$. If $\lambda_t < 1$, then $\Sigma_i(\lambda_t b_t X_{it} +1)^{1/\lambda_t}/I$ is an underestimate of *mean* rent; $\lambda_t = 0$, $\Sigma_i(\lambda_t b_t X_{it} + 1)^{1/\lambda_t}/I$ is an unbiased estimate of *median* rent if the u_{jt} is distributed normally. When we substitute $(\lambda_t b_t X_{it} +1)^{(1-\lambda_t)}$ for $(\lambda_t b_t X_{it} + 1 + \lambda_t u_{it})^{(1-\lambda_t)}$ we are omitting the error terms u_{it}. Although the direct summation of the u_{it} over the I is zero, the same will not be true of the sums raised to a power greater than 1 because of Jensen's inequality. Thus, the measure of central tendency changes as λ_t changes, with that measure increasing toward the mean as λ_t increases toward 1.

An increase in λ from t to t+1, for example, would increase the measured inflation simply because the second-period measure of central tendency would be closer to the mean rather than the median than would the measure of central tendency in the first period. The potential bias associated with increases in λ must be weighed against the alternative of fixing λ across two cross-sections. The commonly used semi-log case is an extreme example of this, with λ fixed at zero. With λ significantly greater than zero, this assumption introduces specification error into the estimates of rental values, but it is not clear whether it imparts a bias into the measured price change. In the next section, we investigate potential bias inherent in changes in λ by constraining λ_{t+1} to equal λ_t when estimating b_{t+1}. To anticipate, we find that the extent of upward bias associated with an increase in λ across adjacent time periods is small.

Hedonic price indexes. Table 8 presents constant quality, Fisher Ideal, hedonic measures of rental inflation, compared with both the published CPI for rent and the CPI adjusted for nonresponse bias, aging bias, and recall bias. Note that all of these adjustments were fully incorporated into the published CPI by 1995 so that the published CPI and the adjusted CPI are the same for the 1995-97 period. There are two areas of particular interest. First, the hedonic measure gives a long-run, average inflation rate of 6.86 percent over the 1975-95 period; that is considerably higher than the published rate of increase, 5.1 percent. Second, if we incorporate all adjustments eventually adopted for the published CPI into the entire series, the adjusted CPI average inflation, 6.29 percent, is considerably closer to the hedonic measure of inflation, 6.86 percent annually, as shown in Table 9. Comparison of the hedonic, published, and adjusted CPIs raises several questions. Is the adjusted CPI measure still too low? Is the pattern of adjustment consistent with the evidence from the hedonic measure? And finally, is the aging adjustment used in the CPI consistent with the estimates underlying the hedonic index?

The finding that the hedonic-based rental inflation estimates exceed the adjusted rental inflation rates by 0.57 percentage point annually raises the issue of whether the adjustments are too small or the hedonic estimates are too high. The average rate of rental-price increase in the adjusted CPI series is essentially the same as that of median gross rents over the sample period. If quality of rental unit were constant over the sample period, this would suggest that the adjusted CPI might be closer to the true rate of rental-price increase than the hedonic measure. Virtually all measures of rental unit quality, however, except average age, increased over the sample period. If quality is increasing, then one would expect quality-adjusted rental prices to appreciate more rapidly than gross rents. This suggests the adjusted CPI series likely understates the rate of rental price increase.

The hedonic rental-price index, on the other hand, potentially has an upward bias associated with the systematic increase of λ over the sample period. To investigate the magnitude of this bias, we compare our Box-Cox-based hedonic estimates with two alternatives: one in which λ is held constant across pairs of cross-sections and one based on the traditional semi-log specification.[17] The three indexes are shown in Table 10. Table 10 shows that the three hedonic indexes all yield similar average rates of rental inflation, although the Box-Cox estimation in which λ varies across cross-sectional pairs does result in slightly higher rates of rental growth, as the potential upward bias would suggest. Holding λ constant reduces the estimated average rate of rental price increase from 6.86 percent annually to 6.73 percent annually. The near identical averages of the λ-constant Box-Cox and the semi-log hedonic indexes suggest that the long-run impact of specification biases associated with the semi-log index is not important. Moreover, there are only modest differences in the patterns of yearly increases across the three indexes.

The only adjustment to the CPI that is clearly reflected in the hedonic index in Table 9 is the 1978 change to eliminate nonresponse bias (which also introduced the recall bias.) Prior to the elimination of the nonresponse bias in 1978 (and the introduction of recall bias), the published CPI was substantially below the hedonic measure – by 3.3 percentage points. After the correction, the difference between the published CPI and our hedonic estimate averaged 1.4 percentage points for the rest of the sample, with no clear pattern in the divergences between the published CPI and the hedonic estimates. Thus there is no clear impact of the 1983 adjustment to eliminate vacancy bias, the 1988 adjustment for aging, or the 1994 adjustment to eliminate recall

[17]Note that λ is held constant by jointly estimating λ_t and β_t then transforming rent in the subsequent cross-section by the estimated λ_t, and estimating β_{t+n} (Laspeyres index, λ_{t+n} , is estimated for the Paasche index).

bias.

Reconciling these changes with the hedonic index is confounded, in part, because the estimations underlying the hedonic indexes do not imply a constant adjustment for aging. Recall that the BLS introduced a constant aging adjustment of 0.36 percentage point in 1988. Our estimates indicate that the impact of aging has been systematically declining over time. In 1975, adding a year to a unit reduced its rent by 1.06 percent. By 1995, that figure had declined to less than a quarter of its 1975 impact and stood at 0.22 percent. Thus in the early years, the aging process introduced much larger downward biases in rental-price indexes than in later years. However, the fact that the BLS currently overstates the aging impact brings the published CPI more in line with the hedonic estimates.

IV. Summary

We have argued in this paper that the rate of rental inflation was quite substantially underestimated in the period from 1942 to 1977 and that in the period since then, this bias has been reduced considerably, although quite probably not eliminated.

We use two techniques to establish this conclusion. First, we model the impact of nonresponse bias and calibrate that model with data from a published study of BLS microdata from the period 1979-81. Second, we use an independent source, the American Housing Survey, to construct a biennial rental inflation measure from 1975 to 1995. Both these measures suggest that prior to 1977, the bias was greater than afterward.

Important questions remain. One question is whether the BLS correction for vacancy bias essentially eliminated the problem beginning in 1983. If the BLS correction eliminated the problem, this could explain much of the discrepancy between our nonresponse bias results and our hedonic measures of rental inflation. Another question is whether hedonic regressions applied to Census of Housing microdata from 1940 to 1980 confirm the basic outlines of the nonresponse bias for that period. A third question is why the hedonic estimates differ so substantially from the CPI measures in the period from 1989 to 1995, when we believe the BLS had, for the most part, eliminated biases.

Table 1 Change in median gross rents compared with CPI for rent					
	Change in median gross rent	CPI for rent	Change in median gross rent minus CPI for rent	CPI for rent, adjusted by authors to current BLS methods	Change in median gross rent minus adjusted CPI for rent
1930-40	-2.3 *	-2.7	0.4	-2.3	0
1940-50	5.2	2.3	2.9	4.2	1
1950-60	5.3	2.7	2.6	4.4	0.9
1960-70	4.3	1.9	2.4	3.1	1.2
1970-75	7.6	4.5	3.1	7.1	0.5
1975-77	8.6	5.9	2.7	9.2	-0.6
1977-83:	9.4	7.4	2	9.3	0.1
1983-87	6.1	5.3	0.8	6.1	0
1987-95	3.4	3.1	0.3	3.7	-0.3

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

Table 2 Data on six-month rent increases Change in rent from one survey to the next (Units are surveyed at six-month intervals) Data collected October 1979 to March 1981, reflecting six-month changes from the period April 1979 to March 1981						
Occupancy status	Number surveyed	Number with six-month rent change	Proportion with rent change	Average rent change for units with change	Average rent change for all units	Average rent change for all units, annualized
6 months or more	37144	17243	46.4 %	8.94	4.15	8.5
5 months or less	8614	6939	80.6 %	12.07	9.72	20.4
all occupants	45758	24182	52.8 %	9.84	5.20	10.7
vacancies (data imputed)	3833	3833	100 %	10.8*	10.8	22.8
occupants and imputed vacancies	49591	28015	56.4 %	9.97	5.51	11.3

Data computed from Rivers and Sommers, pp. 202-203, tables "Analysis of Six-Month Rent Changes by Length of Occupancy." Note: rental information is collected from a rental unit every six months.

*We assume the six-month rent change for vacancies is the same as the change for units with a one-month occupancy.

Table 3 Data on one-month rent increases How much has the rent increased since the last month? Data collected October 1979 to March 1981, reflecting one-month changes from the period September 1979 to March 1981						
Occupancy status	Number surveyed	Number with one-month rent change	Proportion with rent change	Average one-month rent change for units with change	Average % one-month rent change for all units	Annualized % average rent change for all units
6 months or more	37144	1704	4.6 %	10.22*	0.469	5.8
5 months or less	8614	837	9.7 %	11.76** (estimate)	1.141	14.5
all occupants	45758	2541	5.6 %	10.73	0.601	7.5

Source: Rivers and Sommers (1983), pp. 202-203.
*The weighted average of the one-month changes for tenants with six months' or more occupancy by type of unit as reported in the third panel of the table on p. 202 of Rivers and Sommers.
**Average of the one-month change for tenants with six months' or more of occupancy and the six-month change for units with occupancy of one month as reported in the table on p. 203 of Rivers and Sommers.

Table 4 Changes in BLS procedures for collecting rents			
Date	Change	Immediate impact in percentage points per year	General impact for inflation rate π
1952	Reduced rent price collections from quarterly to semiannually	.3	$.075\pi$
1978	Increased response rate for new renters	1.4	$.24\ \pi$
1978	Recall bias introduced	- .6	$-.09\ \pi$
1983	Vacancy imputation	0.9	$.09\ \pi$
1988	Aging bias	0.36	0.36 %
1994	Recall bias	0.22	$.09\ \pi$
Total increase			$.405\ \pi$ +0.36 %

Table 5
Sample Means and Standard Deviations

Variable	1975 Mean	1975 SD	1985 Mean	1985 SD	1995 Mean	1995 SD
Rent	134.39	74.78	314.09	166.40	494.95	235.29
Building characteristics:						
Detached dummy	0.28	0.45	0.24	0.43	0.23	0.42
Multi-unit dummy	0.64	0.48	0.68	0.47	0.67	0.47
Low-rise dummy	0.06	0.24	0.11	0.32	0.11	0.32
Mid-rise dummy	0.02	0.15	NA	NA	NA	NA
High-rise dummy	0.02	0.15	NA	NA	0.06	0.23
Building age	26.53	15.15	30.96	22.34	38.40	23.24
Building age squared	933.15	738.26	1457.13	1692.74	2015.07	2098.45
Garage dummy	NA	NA	0.28	0.45	0.31	0.46
Public sewer dummy	0.85	0.36	0.90	0.30	0.92	0.27
Unit characteristics:						
Number of rooms	4.07	1.41	4.29	1.44	4.35	1.42
Number of bathrooms	1.06	0.38	1.16	0.41	1.21	0.43
Bathroom missing dummy	0.05	0.21	0.02	0.13	0.01	0.09
Central air dummy	0.15	0.35	0.27	0.44	0.33	0.47
Satisfaction with unit	7.24	2.18	7.32	2.41	7.46	2.17
Holes in floor dummy	0.04	0.19	0.03	0.17	0.02	0.13
Mice dummy	0.12	0.32	0.08	0.26	0.05	0.21
Neighborhood characteristics						
Crime dummy	0.07	0.25	0.07	0.26	0.13	0.34
Noise dummy	0.05	0.23	0.11	0.31	0.12	0.33
Trash dummy	0.04	0.20	0.04	0.20	0.04	0.20
Satisfaction with neighborhood	7.53	2.06	7.29	2.69	7.26	2.53
Geography						
Center city location dummy	0.34	0.47	0.48	0.50	0.47	0.50
Within SMSA dummy	0.73	0.45	0.85	0.36	0.87	0.34
Midwest dummy	0.23	0.42	0.23	0.42	0.24	0.42
South dummy	0.30	0.46	0.30	0.46	0.27	0.45
West dummy	0.21	0.40	0.23	0.42	0.27	0.44
Rental terms						
Apartment furnished dummy	0.16	0.37	NA	NA	NA	NA
Rent subsidized dummy	0.02	0.13	0.04	0.20	0.05	0.21
Public housing project dummy	0.07	0.26	0.07	0.26	0.07	0.25
Number of observations	17,207		12,449		15,342	

Table 6
Box-Cox Hedonic Estimations

Dependent Variable: $\frac{(\text{Rent}_{it}^{\lambda} - 1)}{\lambda_t}$	1975[1]		1985[2]		1995	
	Coef	S.E.	Coef	S.E.	Coef	S.E.
Intercept	9.230*	0.173	20.934*	0.700	35.378*	1.401
Lambda	0.38*		0.53*		0.61*	
Building characteristics						
Detached dummy	-0.202*	0.079	0.936*	0.328	2.179*	0.552
Multi-unit dummy	0.739*	0.074	3.588*	0.296	4.440*	0.495
Low-rise	0.884*	0.083	1.955*	0.272	5.342*	0.482
Mid-rise	1.459*	0.128	NA	NA	NA	NA
High-rise	1.744*	0.131	NA	NA	8.355*	0.640
Building age (x100)	-0.042*	0.007	-0.037*	0.016	-0.162*	0.027
Building age squared (x10,000)	-0.0005*	0.0001	0.0007*	0.0002	0.0008*	0.0003
Basement in building	0.371*	0.051	NA	NA	NA	NA
Garage dummy	NA	NA	2.896*	0.200	3.650*	0.344
Public sewer dummy	0.751*	0.061	3.013*	0.290	4.212*	0.593
Condominium	NA	NA	3.890*	0.455	5.137*	0.707
Unit characteristics						
Number of rooms	0.449*	0.017	0.922*	0.069	2.413*	0.121
Number of bathrooms	2.180*	0.070	6.282*	0.253	11.213*	0.397
Bathrooms var. missing	-0.354*	0.116	1.708*	0.650	3.308*	1.513
Central air dummy	1.394*	0.066	2.717*	0.219	3.170*	0.360
Satisfaction with unit	0.009	0.011	-0.197*	0.042	0.004	0.080
Holes in floor dummy	-0.057	0.103	-2.108*	0.483	-0.511	1.035
Mice dummy	-0.466*	0.062	-1.077*	0.318	-0.584	0.653

Neighborhood characteristics						
Crime dummy	0.241*	0.083	0.677*	0.326	-0.328	0.475
Crime variable missing dummy	NA	NA	NA	NA	NA	NA
Abandoned buildings nearby	-0.570*	0.068	-2.686*	0.354	-6.553*	0.641
Noise dummy	0.199*	0.087	-0.078	0.275	-0.047	0.465
Trash dummy	0.120	0.102	0.289	0.403	-1.281	0.706
Satis. with neighborhood	0.085*	0.012	0.186*	0.040	0.385*	0.074
Location						
Center city location	-0.271*	0.047	-1.386*	0.182	-2.346*	0.312
Within SMSA	1.440*	0.051	6.274*	0.248	13.143*	0.451
Midwest dummy	-1.147*	0.056	-5.261*	0.241	-11.072*	0.423
South dummy	-2.131*	0.064	-6.290*	0.249	-14.613*	0.464
West dummy	-0.436*	0.066	-0.086	0.261	-0.361	0.459
Rental terms						
Apartment furnished	0.508*	0.056	NA	NA	NA	NA
Rent subsidized	-2.497*	0.144	-9.635*	0.390	-3.366*	0.646
Public housing project	-3.636*	0.077	-7.747*	2.116	-16.882*	2.619
Adjusted R^2	0.538		0.450		0.401	
Number of observations	17,207		12,448		15,341	

* Denotes significance at the 5% level

Table 7
Estimated λ, 1975-1995

Year	λ
1975	0.38
1977	0.40
1979	0.40
1981	0.40
1983	0.42
1985	0.54
1987	0.52
1989	0.56
1991	0.53
1993	0.58
1995	0.61

Table 8
AHS median gross rents and constant quality hedonic measure of rental inflation compared to published and adjusted CPI for rent

	Changes in median gross rent	Hedonic measure	CPI, rent, IVQ to IVQ	Adjusted CPI, rent, IVQ to IVQ	Adjustment
1975-77	8.6	9.3	5.9	9.2	3.3
1977-79:	8.6	9.0	7.7	9.6	1.9
1979-81	11.5	11.2	8.7	10.8	2.1
1981-83	8.0	7.0	5.9	7.5	1.6
1983-85	7.5	8.7	6.0	7.0	1
1985-87	4.7	5.5	4.5	5.3	0.8
1987-89	3.1	5.4	3.9	4.3	0.4
1989-91	4.4	5.9	3.6	3.9	0.3
1991-93	2.7	2.7	2.3	2.5	0.2
1993-95	3.6	3.8	2.5	2.8	0.3
1995-97	2.5	NA	2.9	2.9	0
Average, 1975-95	6.27	6.86	5.1	6.29	1.19

Sources: American Housing Survey, CPI, and authors' calculations.

Table 9
Constant quality, Box-Cox, measure of rental inflation compared to adjusted CPI for rent

	Hedonic measure	Adjusted CPI, rent, IVQ to IVQ	Difference
1975-77	9.3	9.2	-0.1
1977-79:	9.0	9.6	0.6
1979-81	11.2	10.8	-0.4
1981-83	7.0	7.5	0.5
1983-85	8.7	7.0	-1.7
1985-87	5.5	5.3	-0.2
1987-89	5.4	4.3	-1.1
1989-91	5.9	3.9	-2
1991-93	2.7	2.5	-0.2
1993-95	3.8	2.8	-1
Average, 1975-95	6.86	6.29	-0.57
Average, 1985-95	4.66	3.76	-0.9

Sources: American Housing Survey, CPI, and authors' calculations.

Table 10
Alternative Hedonic Estimates
(Annual Percent Growth)

Year	Box-Cox	Box-Cox λ-constant	Semi-log
1975-77	9.3	9.2	9.2
1977-79	9.0	9.0	8.9
1979-81	11.2	11.2	11.0
1981-83	7.0	6.9	6.7
1983-85	8.7	8.0	8.5
1985-87	5.5	5.6	5.3
1987-89	5.4	5.2	4.9
1989-91	5.9	6.1	6.6
1991-93	2.7	2.5	2.7
1993-95	3.8	3.7	3.7
Average	6.86	6.74	6.74

Figure 1.a
Actual Rent, Semi-log, and Box-Cox Predicted Rent

1975

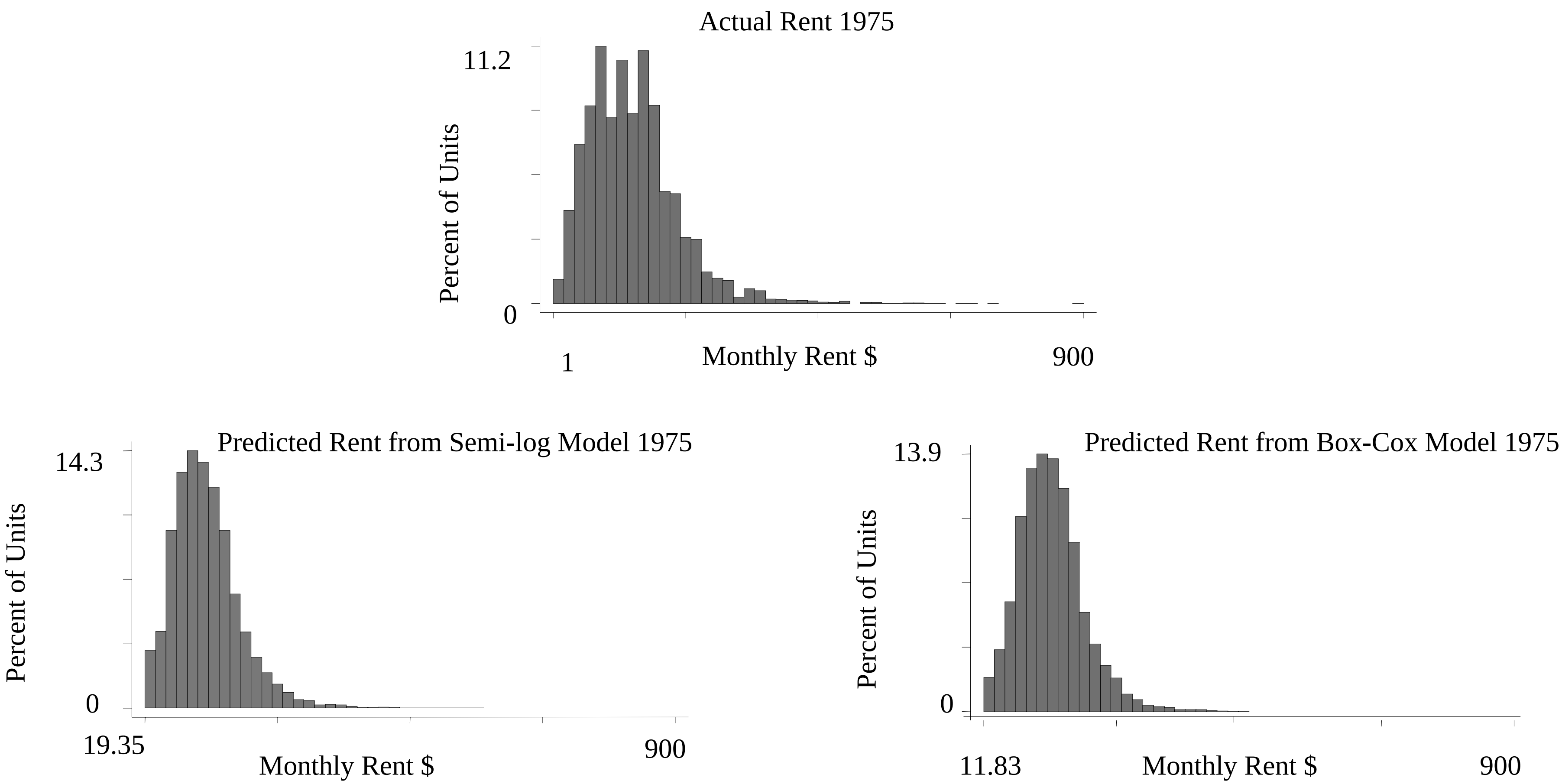
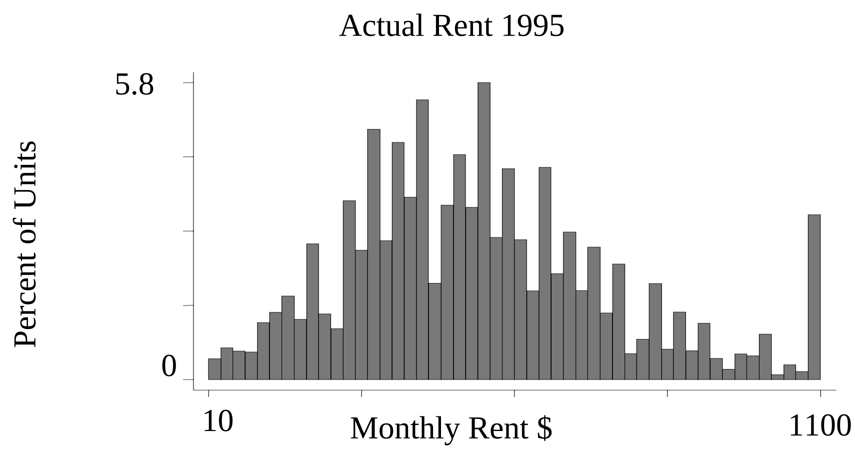
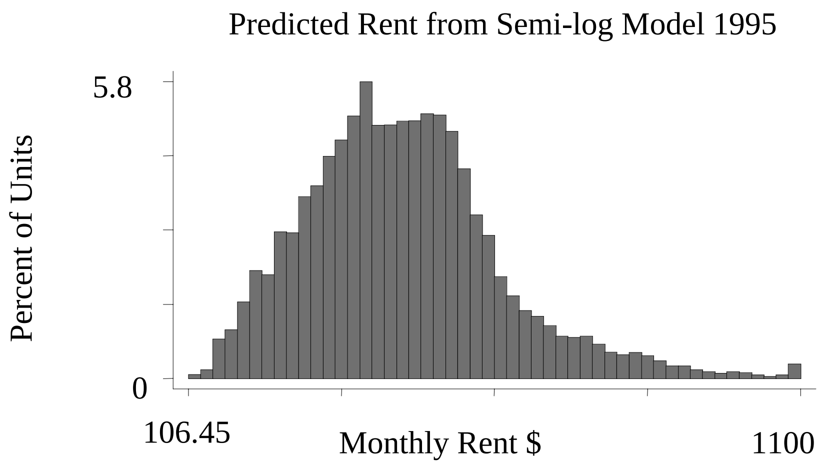


Figure 1.b
Actual Rent, Semi-log, and Box-Cox Predicted Rent
1995





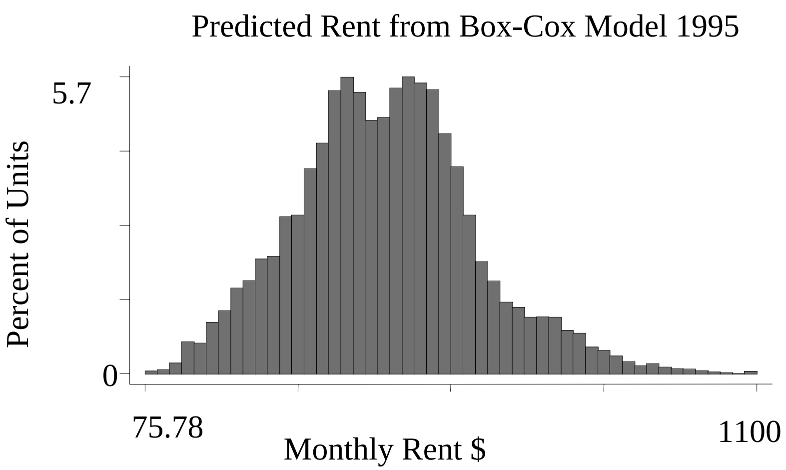
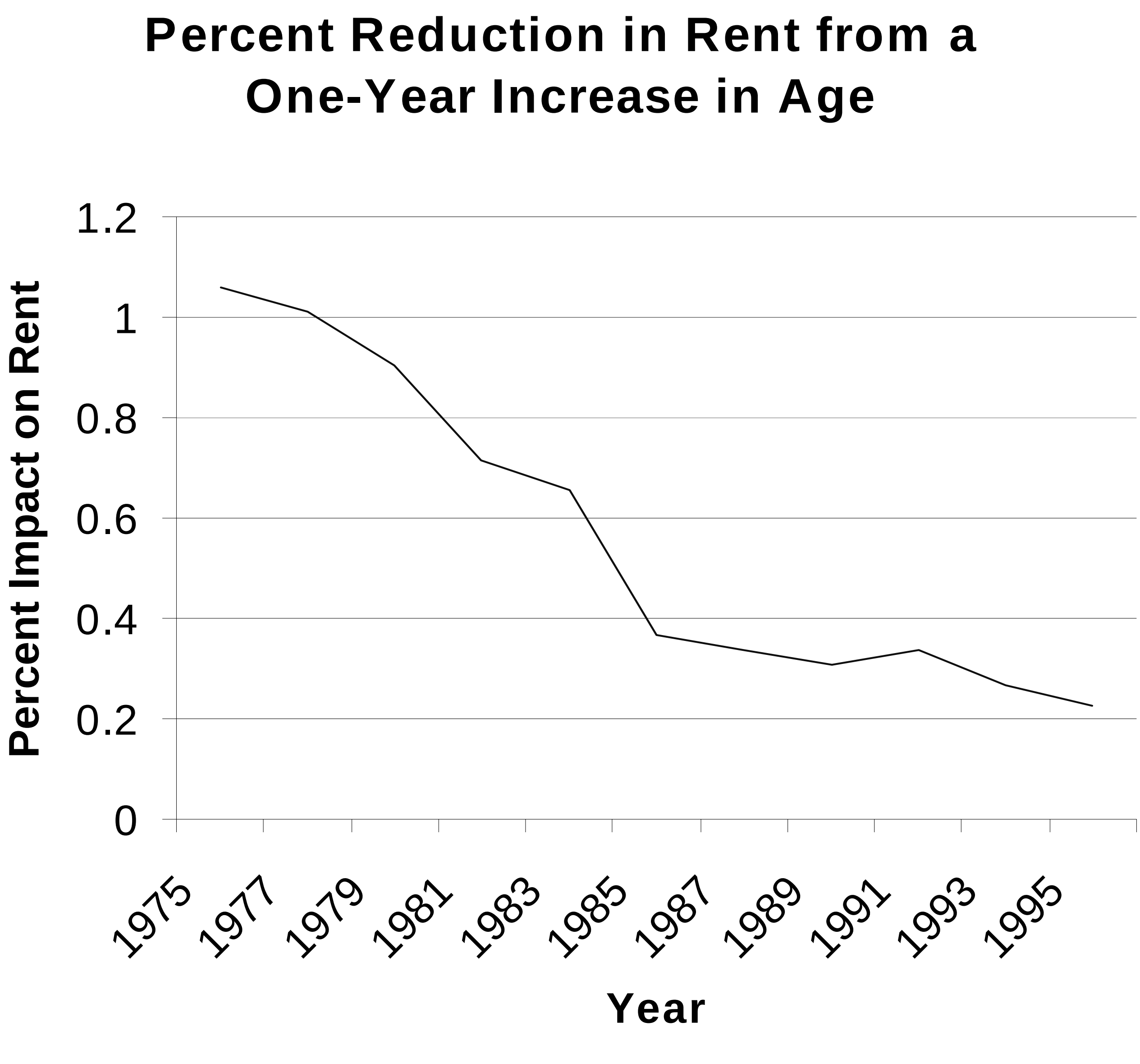


Figure 2


Percent Reduction in Rent from a One-Year Increase in Age
Percent Impact on Rent
1.2
1
0.8
0.6
0.4
0.2
0
1975
1977
1979
1981
1983
1985
1987
1989
1991
1993
1995
Year

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