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Wages and prices: Are workers fully compensated for cost of living differences?☆

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ABSTRACT

This paper investigates the equilibrium relationship between wages and prices across labor markets. Of central interest is the extent to which workers receive higher wages to compensate for differences in the cost of living. According to the spatial equilibrium hypothesis, the utility of homogenous workers should be equal across labor markets. This implies that controlling for amenity differences across areas, the elasticity between wages and the general price level across areas should equal one, at least under certain conditions. I test this hypothesis and find that the predicted relationship holds when housing prices are measured by rents and the general price level is instrumented to account for measurement error. When housing prices are measured by housing values, however, the wage-price elasticity is significantly less than one, even using instrumental variables. Rents reflect the price paid for housing per unit of time and are arguably the superior measure. Thus, findings in this essay provide support for the full compensation hypothesis. These findings also have important implications for researchers estimating the implicit prices of amenities or ranking the quality of life across areas.

1. Introduction

A number of studies have shown that wages differ across labor markets even after control for observable individual characteristics.[1] Such wage dispersion across markets can in part be attributed to differences in prices and amenities across areas. If a city has higher prices for goods and services providing a given level of utility, workers will require higher wages to work there.[2] Similarly, if a city has nicer amenities, all else the same, workers will be willing to accept lower wages to work there. In order for a spatial equilibrium to occur, utility must be equal across areas for workers with identical skills and preferences. In previous literature, this is sometimes referred to as the competitive hypothesis or the law of one wage. Many studies have attempted to test the competitive hypothesis (e.g. regional wage gap studies), but they are often hindered by limited information on area prices and amenities.

Several studies interested in interarea wage differentials have used an interarea price index to fully adjust wages for price differences by dividing nominal wages by the price index.[3] Other studies have used fully-adjusted wages to measure the implicit prices of amenities across cities (e.g. Rosen, 1979; Greenwood et al., 1991; Glaeser and Tobio, 2008).[4] DuMond et al. (1999), however, suggest that full adjustment for prices may be inappropriate to measure interarea wage differentials. They instead advocate using a partial adjustment whereby the log of the price index (and potentially higher order terms) is included as an independent variable in a log wage equation. The coefficient on the log of the price index can be interpreted as the wage-price elasticity. One hypothesis is that the elasticity between wages and the general price level is equal to one.[5] I refer to this as the full compensation hypothesis. Researchers who fully adjust wages for prices implicitly assume that the full compensation hypothesis holds, but few studies have explicitly tested the full compensation hypothesis.

Two studies that have estimated the elasticity between wages and prices are Roback (1988) and DuMond et al. (1999). Roback (1988) uses a now discontinued cost of living index produced by the Bureau of Labor Statistics and estimates a wage-price elasticity of 0.97, both with and without controls for amenities, which would seem to lend support for the full compensation hypothesis. As discussed below, a reexamination of Roback (1988), however, suggests that her measurement of prices is inappropriate and biases her estimates. DuMond

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[1] See Dickie and Gerking (1989) for an early review of the literature on interarea wage differentials in the United States.

[2] In this paper, I often use the term city to refer to metropolitan areas.

[3] See for example, Coelho and Ghali (1971, 1973), Bellante (1979), Gerking and Weirick (1983), Johnson (1983), Sahling and Smith (1983), Dickie and Gerking (1987), and Farber and Newman (1987).

[4] See Gyourko et al. (1999) for a review of the literature on amenity valuation and quality of life.

[5] Throughout this paper, I often refer to the "general price level" in a city. I mean by this a composite price index based on a fixed basket of housing and non-housing goods weighted by the consumption shares of the relative goods.

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et al. (1999) use a price index based on the ACCRA *Cost of Living Index* and find a wage-price elasticity of 0.46 controlling for amenities and 0.37 absent amenities. Thus, the magnitude of the wage-price elasticity and validity of the full compensation hypothesis are still open questions.

This paper builds on earlier work by examining the *equilibrium* relationship between wages and prices, controlling for amenities. I stress the word *equilibrium* because wages and prices are simultaneously determined. While this paper does not provide evidence on the causal effect of prices on wages or *vice versa*, much can be learned from examining the equilibrium relationship between the two. Following Rosen (1979)and Roback (1982), I develop a model that predicts that under certain conditions (e.g. Cobb–Douglas utility) the elasticity between wages and the general price level should equal one controlling for amenities. In other words, workers should be fully compensated for differences in prices across cities. However, to the extent that the assumptions of the model do not hold, the elasticity between wages and the general price level may differ from unity. The relationship between wages and prices is ultimately an empirical question.

I find that estimates of the wage-price elasticity are sensitive to whether housing prices are measured by housing values or rental payments. Rents are the ideal measure of housing prices, the price paid per unit of time for the use of housing, but in practice housing values are often used to measure housing prices. My preferred specification measures housing prices by rents. Measuring housing prices by rents and using Ordinary Least Squares, I estimate the wage-price elasticity to equal 0.76, but OLS estimates may be downwardly biased due to measurement error in the price index, especially the non-housing price component. Instrumenting for the rent-based price index using rents for the previous year, the estimated elasticity between wages and the general price level is nearly identical to one. Again, if rents are the ideal measure of housing prices, this finding provides strong empirical support for the full compensation hypothesis.

When housing prices are measured by housing values, the estimated elasticity between wages and the general price level is never more than 0.5, even using instrumental variables. The findings of this paper have important implications for researchers estimating the implicit prices of amenities or ranking the quality of life across areas. First, when adjusting wages for prices, housing prices should be measured by rents and not values. Second, it is shown that ignoring differences in non-housing prices, as often done, biases estimates of the implicit prices of amenities.

2. Theoretical considerations

This section develops a simple model of the equilibrium relationship between wages, prices, and amenities across cities and regions following Rosen (1979) and Roback (1982). Firms produce X_1 and X_2 according to constant returns to scale production functions using labor (N), capital (K), and land (L) given locational differences in productivity due to amenities (Z): $X_i = X_i(N, K, L; Z)$. The marginal products of labor, capital, and land are all non-negative, but increases in amenities can either increase or decrease productivity. The price of capital is determined exogenously in the world market and normalized to equal one, while the prices of labor (W) and land (P_L) are determined competitively in local markets. In equilibrium, firms earn zero profits and the price of each good is equal to its unit cost of production (C_i):

$$C_i(W, P_L; Z) = P_i, \quad i = 1, 2. \tag{1}$$

Workers maximize utility subject to a budget constraint, where utility is a function of goods X_1 and X_2 and location-specific amenities: $U = U(X_1, X_2; Z)$. Workers are mobile across cities and regions, and in equilibrium utility for identical workers is equal across areas. The indirect utility function can be represented as a function of wages and the prices of X_1 and X_2 given amenities:

$$V = V(W, P_1, P_2; Z). \tag{2}$$

Taking the total differential of both sides of Eq. (2), setting $\mathrm{d}V = 0$, rearranging, and employing Roy's Identity yields a slight variant of the equation used by Roback to estimate the implicit price of amenities (Eq. (5) in Roback, 1982):

$$\mathrm{d}W = X_1\mathrm{d}P_1 + X_2\mathrm{d}P_2 - P_Z\mathrm{d}Z.^{6} \tag{3}$$

However, instead of solving for the price of amenities (P_Z), the equation is solved for dW. Dividing both sides of Eq. (3) by W, converts the equation to logarithmic form:

$$\mathrm{d}\ln W = (P_1X_1 / W)\mathrm{d}\ln P_1 + (P_2X_2 / W)\mathrm{d}\ln P_2 - (P_Z / W)\mathrm{d}Z. \tag{4}$$

Eq. (4) says that controlling for amenities, a one percent increase in the price of X_1 will require wages to increase by a percentage equal to the share of wages spent on X_1 in order for utility to remain constant. The same is true for increases in the price of X_2, and the result easily generalizes to the case of more than two goods. In other words, the wage-price elasticity for a good should be equal to the budget share of the good, assuming that non-wage income is negligible. Furthermore, if total consumption expenditure is equal to wage income, $P_1X_1 + P_2X_2 = W$, then a one percent increase in the prices of all goods will require wages to increase by one percent to maintain equal utility.

While this interpretation of Eq. (4) is valid for small changes in prices, it may be less valid for large changes in prices as consumers respond to large price differences by altering their consumption mix. However, if utility is Cobb–Douglas as suggested by Davis and Ortalo-Magné (2008) and others, the elasticity between wages and the price of a good is equal to the expenditure of a good even for large changes in prices. To see this, let utility take the Cobb–Douglas form: $U = f(Z)X_1^{\alpha}X_2^{(1-\alpha)}$. Taking a monotonic transformation the indirect utility function can be written as: $V = C + \ln W - \alpha\ln P_1 - (1-\alpha)\ln P_2 + \ln(f(Z))$, where α is the constant budget share for X_1, $(1-\alpha)$ is the budget share for X_2, and C is a constant. Holding utility constant across areas, $\partial\ln W / \partial\ln P_1$ is equal to α even for large changes in prices. In other words, Cobb–Douglas utility suggests that the elasticity between wages and the price of a good is equal to the good's budget share even for large price changes. Similarly, Cobb–Douglas utility predicts that the elasticity between wages and the general price level should equal one. Workers would, therefore, require full compensation for price differences across cities.

The full compensation hypothesis has considerable intuitive appeal. Suppose there are two cities with equal bundles of consumer amenities, but one city has higher prices for goods and services. If the general price level in the expensive city is 10% higher than in the less expensive city, how much higher will wages have to be in the expensive city to keep workers from leaving for the other city? Intuition seems to suggest that a worker would need 10% higher wages to compensate for the 10% higher price level. In other words, workers would require full compensation for price differences holding amenities constant.

Workers may not be fully compensated for price differences for a number of reasons. If workers are highly immobile or do not have sufficiently good information on wages, prices, and amenities in other cities, then migration may not arbitrage away interarea differences in wages, prices, and amenities. In other words, barriers to migration may cause workers in some markets to have higher utility levels than comparable workers in other markets. In reality though, workers are

[6] Alternatively, we could have defined the expenditure function and used Shephard's Lemma to obtain an equivalent result as in Albouy (2008b).

often quite mobile across markets. Even if some are relatively immobile, the movement of marginal migrants between labor markets may result in an equilibrium relationship between wages and prices that yields equal utility across areas for all homogenous workers.

The relationship between wages and prices may also differ from full compensation if utility is considerably different from Cobb–Douglas and prices are very different across markets. Thinking of X_1 and X_2 in the above model as housing and non-housing consumption, a high degree of substitutability between housing and non-housing may cause the true elasticity between wages and the general price level to be less than one.[7] As will be shown later, housing prices are significantly more dispersed across areas than non-housing prices. If workers can easily substitute away from housing consumption in places where it is relatively expensive, they will not have to be fully compensated for differences in housing prices.[8] As a result, a fixed basket price index will overstate the true cost of living in expensive cities and cause the elasticity between wages and the general price level to be less than one.

As hinted above, the wage-price elasticity also depends on the extent to which people save. If consumption is less than wage income ($P_1X_1+P_2X_2<W$), the true wage-price elasticity should be less than one. Conversely, if consumption is greater than wage income, the wage-price elasticity may be greater than one. Evidence from the 2005 Consumer Expenditure Survey suggests that average consumer expenditures are indeed quite close to average after-tax wage income. The ratio of average expenditures to average after-tax income in the 2005 CES is 0.94. The CES is a relatively small sample and there could be some misreporting (e.g. of income), but the available evidence indicates that assuming expenditures are equal to wage income may be a reasonable first approximation.

There are, therefore, a number of reasons why the elasticity between wages and the general price level may be less than one. Ultimately, the relationship between wages and prices is an empirical question. I explore this relationship empirically in subsequent sections.

3. Empirical considerations/previous literature

The theoretical model suggests that under certain conditions, the elasticity between wages and a composite price index is approximately one. Based on the intuition behind this result, a number of researchers interested in interarea wage differentials have fully-adjusted nominal earnings using an interarea price index and estimated log wage equations of the form:

$$\ln\left(W_{ij}/P_j\right) = X_{ij}\beta + \varepsilon_{ij}, \tag{5}$$

where W_{ij} is the wage for person i in city j, P_j is the price level in city j, X is a vector of personal characteristics, β is the corresponding coefficient vector, and ε is an error term with mean equal to zero. Along these lines, Johnson (1983) obtains the seemingly surprising result that fully-adjusted wages are more dispersed across cities than nominal wages, at least for men.[9] DuMond et al. (1999), however, argue that full adjustment may be inappropriate. Instead, they advocate using a partial adjustment where the dependent variable is the log of the nominal wage and the log of the price index is included as an independent variable on the right hand side:

$$\ln W_{ij} = X_{ij}\beta + \theta\ln P_j + \varepsilon_{ij}. \tag{6}$$

Doing so, they find wage dispersion to be considerably lower across markets than with either nominal or fully-adjusted wages.[10]

Theory and empirics also suggest that wages are affected by attributes that make a city a more or less pleasant place to live. Therefore, Eqs. (5) and (6) can also be modified to include city-specific amenity levels and a corresponding coefficient vector. The parameter θ in Eq. (6) can be interpreted as the interarea wage-price elasticity. If $\theta=1$, Eqs. (5) and (6) are equivalent. However, if θ is not equal to one, Eq. (5) may be misspecified. Thus the value of θ is of considerable interest.

Roback (1988) estimates Eq. (6) both with and without amenities and produces estimates of θ equal to 0.97 for both specifications. DuMond et al. (1999), however, estimate a point estimate for θ of 0.46 with amenities and 0.37 without amenities with standard errors small enough for both to easily reject the hypothesis that $\theta=1$. There are several differences between the two studies, such as the time period considered, the number of cities considered and the amenities included. However, the most important difference is likely the price indices used and the way they are used. Roback uses a now discontinued price index produced by the Bureau of Labor Statistics from the Handbook of Labor Statistics, and DuMond et al. use a price index based on the ACCRA *Cost of Living Index*. Measurement error may be more significant in the ACCRA index, and this may explain some of the difference between the estimates of Roback (1988) and DuMond et al. (1999). DuMond et al. reestimate their results using the BLS *Urban Family Budget and Comparative Indexes for Selected Urban Areas* updated from its 1981 value (the last year it was produced) using the city-specific CPI for a limited number of cities and find that the estimate of θ without amenities increases to 0.526. This price index is much closer to the index used by Roback, but the coefficient estimate it yields is still much less than one.

Closer examination of the two studies reveals a more subtle distinction in the way the price indices are used. DuMond et al. (1999) use the same price index for all workers within a given city. In Roback (1988), on the other hand, the price variable used consists of "low, medium, and high standard of living budgets assigned based on individual family income and number of dependents" (p.41). In other words, Roback assigns persons within a given city a different budget based on their income. Presumably, her intent is to assign to each individual the most relevant price for their particular consumption bundle. This approach creates intra-city variation in prices, and a problem arises if the intra-city variation in prices is spuriously correlated with intra-city differences in wages. In such a case, the coefficient on the price variable in the log equation will be biased.

The budgets formerly produced by the BLS and used by Roback (1988) are based on what it would cost a family of four in a given city to obtain a given standard of living. The BLS computes the budgets (B_{ij}) for each standard of living (i) in each city (j) by multiplying local prices (P_{ij}) by a basket of goods (X_{ij}) for each standard of living. The basket is also allowed to vary across cities within a standard of living, but is intended to maintain a given standard of living across cities.

[7] Cobb–Douglas utility implies an elasticity of substitution equal to one. The limited literature has not reached a consensus on the elasticity of substitution between housing and non-housing. Ogaki and Reinhart (1998) estimate the elasticity of substitution to be 1.17, but not statistically different from one at the 5% significance level. Piazessi et al. (2007) find estimates of 0.77 and 1.24 depending on the time period considered, neither of which is statistically different from one. However, Benhabib et al. (1991) and McGrattan et al. (1997) estimate the elasticity of substitution to be 2.5 and 1.75, respectively. Davis and Ortalo-Magné (2008) do not explicitly estimate the elasticity of substitution, but do find that the expenditure share on housing is roughly constant over time and across metropolitan areas suggesting that the elasticity of substitution is close to one.

[8] Consumers can also shift away from consumption of relatively expensive housing toward consumption of local amenities, especially since local residents can often consume natural amenities at very low marginal cost (e.g. climate and coastal location). As suggested by an anonymous referee, this appears quite likely along parts of the California coast, where good weather permits substitution of outdoor living for indoor living.

[9] Johnson (1983) uses a pooled cross-section of 34 cities from the May Current Population Survey for 1973–1976 with price data from the BLS for an intermediate standard of living for 1974.

[10] DuMond et al. (1999) use a pooled cross-section of 185 cities from the 1985–1995 CPS Outgoing Rotation Group files with price data from the ACCRA *Cost of Living Index* from the same period.

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Ignoring temporarily that the basket varies across cities, recognize that $B_{ij}=P_{ij}X_i$. Regressing $\ln W_{ij}$ on $\ln(P_{ij}X_i)$ is clearly not the same as regressing $\ln W_{ij}$ on $\ln P_{ij}$ because X_i is increasing with income. If one were to use the same budget, $B_j=P_jX$, (e.g. the intermediate standard of living budget) for all workers within a given city and hence have no intra-city variation in budgets, then there would be no problem because taking logs causes $\ln X$ to drop into the constant term. Using budgets instead of price index values and allowing the budgets to vary across types of workers within cities means that the "price" variable is severely confounded by intra-city variations in consumption. In other words, the estimates are biased by the fact that workers within a city who have higher wages also have higher standards of living and are assigned a higher consumption basket. In work not shown, I attempt to replicate Roback (1988). The results suggest that measuring prices by budgets and allowing intra-city variation in prices does indeed upwardly bias estimates of the wage-price elasticity.[11]

Henderson (1982) also estimates a variant of Eq. (6) that includes the log of housing prices instead of a composite price index. Henderson is also one of the few studies in this area to look at after-tax earnings instead of pre-tax earnings. He finds point estimates of 0.17 and 0.21 for the coefficient on log housing prices in alternate specifications that vary in the amenities included. Henderson does not incorporate non-housing prices in his regressions, however, and he measures housing prices by ownership costs, though rents are likely preferable.

Recent papers by Albouy (2008b) and Davis and Ortalo-Magné (2008) are also interested in the relationship between wages and prices. Albouy (2008b) attempts to construct improved quality of life rankings for cities by among other things incorporating non-housing prices and federal income taxes into the rankings. His main finding is that improved quality of life estimates rank large cities more favorably than has been the case using previous methods. He also computes city fixed effects for log housing prices and log wages and regresses the log housing prices on log wages and amenities. The regression yields a coefficient of 1.41. Based on his chosen parameters (for the budget shares of housing and non-housing, etc.), he suggests that his model quite accurately predicts the relationship between housing prices and wages across cities. However, he measures housing prices using combined data on housing values and rents. The results in the current paper suggest that the relationship between wages and rents is considerably different from the relationship between wages and housing values.

Davis and Ortalo-Magné (2008) develop a model of the equilibrium relationship between wages and prices across cities that assumes a Cobb–Douglas utility function and therefore that the expenditure share for housing is constant across cities. They test their model by predicting city-specific rental values as a function of wages and comparing predicted rents to observed rents, where quality is held constant for both housing and labor. They find that observed rents are under-dispersed compared to what is predicted by their model, i.e., rents are too low in many high wage areas and too high in many low wage areas. Davis and Ortalo-Magné suggest that the omission of amenities from their analysis may partially explain their results.

4. Data and methods

In the empirical section of this paper, I begin by estimating a variant of Eq. (6) that includes amenities (Z):

$$\ln W_{ij} = X_{ij}\beta + \theta \ln P_j + \gamma Z_j + \varepsilon_{ij}. \tag{7}$$

I use earnings and individual characteristics data from the 2006 Current Population Survey Outgoing Rotation Group (CPS-ORG) files merged with data on prices and amenities from several sources.[12] My sample consists of all employed wage and salary workers ages 18–61 (inclusive), who are not full-time students. I also exclude all persons with imputed earnings to avoid imputation bias, which would bias θ toward zero (Hirsch and Schumacher, 2004).[13] The dependent variable is the log of the hourly wage ($\ln W_{ij}$). I use the reported hourly wage for workers who are paid by the hour and do not receive tips, commissions, or overtime. For all other workers, the hourly wage is computed by dividing usual weekly earnings by the usual number of hours worked per week.

My preferred estimates adjust wages for federal income taxes, but I also estimate Eq. (7) using pre-tax wages for the sake of comparison. As discussed by Henderson (1982) and Albouy (2008a,b), the progressivity of the federal income tax system causes workers in high wage areas to pay a higher percentage of their income in federal income taxes than workers in relatively low wages areas. The marginal benefit, however, to an individual worker of her federal income tax contributions is zero because workers consume the same level of federal public services regardless of their federal tax payments. In other words, while workers pay higher federal income taxes in high wage areas, they do not receive higher federal benefits. Consequently, when choosing among cities, workers are concerned with the wages they would earn net of federal taxes in each city instead of gross wages.

The present study does not adjust wages for social security contributions or state income tax payments. It would be relatively straightforward to estimate social security contributions for individual workers, but estimating the benefits to workers of their contributions would be more difficult. I could also estimate state income tax payments for workers, but adjusting wages for state income taxes is inappropriate unless we also adjust wages for other state and local taxes because states differ in their reliance on income taxes. Even if we could compute the total burden of all state and local taxes to each worker, we would still need to account for the benefits from state and local expenditures that each worker receives. Given the complexities involved with estimating the net fiscal incidence of social security payments and state taxes, I make no adjustment for them in the dependent variable.[14] Because the dependent variable in this study is the log of the hourly wage, the analysis is only affected by social security payments and state taxes to the extent that their net fiscal incidence is not proportional to wages for homogenous workers in different areas.[15] However, to the extent that the total net burden of social security and state and local taxes and expenditures for homogenous workers is higher (lower) in high wage areas, regression estimates of θ that only account for federal income taxes may overstate (understate) the true value of θ.

Federal income tax liabilities are not reported in the CPS-ORG files, but are instead estimated using the federal tax schedule and based on several assumptions. I assume that all married couples file jointly and receive two personal exemptions and non-married persons have a filing status of single and receive one personal exemption. Itemized deductions are assumed to equal 20% of annual earnings, where

[11] Using a single price index for each city, I find that the wage-price elasticity using 1973 data ranges between 0.4 and 0.7 depending on which price index is used. However, this replication of Roback (1988) does not include amenities and does not account for measurement error in the price index. In subsequent sections of this paper, I estimate the wage-price elasticity using more recent data controlling for amenities and using instrumental variables to account for measurement error.

[12] Prices and amenities are measured at the city level, where a city is defined as a Core Based Statistical Area or a Combined Statistical Area.

[13] Imputation bias would likely result because imputed earners are often assigned wages of workers in different metropolitan areas or even different regions.

[14] Hence, use of the term after-tax wages implies wages net of federal income taxes only.

[15] To illustrate, suppose we have an equal rate tax (τ) on wages (W) in all areas. Wages net of the tax are $W(1-\tau)$. Because the dependent variable is in logs, note that $\ln W(1-\tau)=\ln W+\ln(1-\tau)$. Because t is a constant, regression results will be equivalent (except for the constant term in the regression) regardless of whether the dependent variable is the log of pre-tax wages ($\ln W$) or the log of after-tax wages ($\ln W(1-\tau)$).

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annual earnings are equal to usual weekly earnings times 48.3 (the average number of weeks worked for workers in the March CPS). Taxpayers take the standard deduction if it is more than their itemized deductions. Deductions and exemptions are subtracted from annual earnings to estimate taxable income. Tax schedules are then used to compute federal tax liabilities. I next compute the average tax rate for each taxpayer (τ_{ij}), and then multiply the hourly wage by one minus the average tax rate to compute after-tax hourly wages ($W_{ij}(1-\tau_{ij})$)).

All regressions include a number of individual characteristic variables intended to make workers roughly similar across cities. The individual characteristics included are eleven dummy variables for highest level of education received, a quartic specification for experience, and dummy variables for race/ethnicity (Black, Asian, Hispanic, and Other), female, married, employed part-time, enrolled part-time in school (measured for workers under 25), union member, naturalized citizen, and non-citizen. Additionally, I include nine occupation dummies, eleven industry dummies, and three dummies for whether the worker is a federal, state, or local government employee.[16] I also include 11 month-in-sample dummies.

The baseline price index is constructed using the ACCRA *Cost of Living Index* for 2006. The ACCRA index is produced quarterly based on prices collected by local chambers of commerce for a basket of 57 goods and services meant to be representative of actual consumer expenditures.[17] The prices of the 57 goods and services are then weighted (based somewhat on CES expenditure data) to form a composite price index and six sub-indices for housing, groceries, utilities, transportation, healthcare, and miscellaneous goods and services.

The baseline price index based solely on ACCRA data, however, may not accurately measure intercity variation in prices. One prominent reason is that ACCRA measures housing prices as a weighted average of the price of two goods: apartment rent and homeowner principal and interest, with homeowner expenses being given a much greater weight (.82) than apartment rent (.18). Housing rents measure the price paid per unit of time for the use of housing, and are therefore the ideal measure of housing prices.[18] Homeowner expenses may be an inappropriate measure of the user cost of housing because they are based on housing values. Homeownership involves both a consumption decision and an investment decision, and the value of a house is equal to the expected net present value of the income stream it generates. If expected future growth in rents differs across cities and over time, then so will the ratios of rents to housing values. Empirical evidence suggests that this is indeed the case (Clark, 1995; Davis et al., 2008). Housing values may even be subject to bubbles based on irrational speculation about the growth in future benefits (Case and Schiller, 2003). Therefore, measuring housing prices using house values is likely to be inappropriate because house values are not based solely on the present user cost of housing.[19] This may be especially true for recent years given the relatively large increase in housing values, especially in several metropolitan areas with a relatively inelastic supply of housing (Glaeser et al., 2008).

Additional difficulties arise with the ACCRA index because prices are not reported for all areas in each year. This has two drawbacks. First, ACCRA often contains no information on prices for a given city, and hence I must exclude the city from the analysis. This limits my analysis to 167 cities, though the cities that remain account for 68% of workers in the CPS. A second problem is that prices are reported at the sub-metropolitan level and must be aggregated to produce city-level averages using population weights, yet not all areas within a metropolitan area are necessarily included. To the extent that sub-metropolitan areas for which prices are reported are not representative of areas in the same city for which prices are not reported, the average price level in the city will be measured with error. Koo et al. (2000) provide a further discussion of issues associated with using the ACCRA index to measure interarea price differences.

To address the potential problems that result from using ACCRA data to measure housing prices, I also compute a modified price index that measures housing prices solely by rental costs from the 2006 American Community Survey (ACS).[20] To do this, I use microdata available from the Integrated Public Use Microdata Series (IPUMS) produced and distributed by Ruggles et al. (2008) to estimate quality-adjusted average gross rents for each city in the sample.[21] The first step is to regress log gross rents, R, for each housing unit on a vector of housing characteristics, F, and a vector of city-specific fixed effects, α:

$$\ln R_{ij} = F_{ij}\Gamma + \alpha_j + u_{ij}. \tag{8}$$

The housing characteristics included are dummy variables for the number of bedrooms, the total number of rooms, the age of the structure, the number of units in the building, modern plumbing, modern kitchen facilities, and lot size for single-family homes. The results from this estimation are available upon request. I then use the estimated parameters to predict average gross rents for each city holding the housing characteristics constant at their mean level for the entire sample.[22] I then divide the quality-adjusted average gross rents for each city by the mean across cities and multiply by 100 to create a housing price index based on quality-adjusted gross rents. I then compute a modified composite price index by taking a weighted average of the rent-based housing price index and non-housing prices from ACCRA, where housing prices are given a weight of 0.29 and non-housing prices are a given a weight of 0.71.[23] Weights are chosen based on calculations from the 2005 Consumer Expenditure Survey suggesting that housing (based on gross rents) represents 29% of average consumption expenditures.[24]

[16] Some individual characteristics are still unobserved. If individuals with high unobserved ability sort into expensive cities, the estimated wage-price elasticity may be positively biased.

[17] While many of the goods in the index might be thought of as traded goods, the law of one price does not strictly hold because most goods are sold at retail. Retailing in San Francisco is more expensive than retailing in Topeka, KS because of higher commercial land rents and higher wages needed to compensate for higher housing rents (and subsequently higher non-housing costs). The spread of online shopping is likely to have important effects in pushing homogenous goods towards a single price, but this is not accounted for under current ACCRA methods.

[18] For this reason, the Consumer Price Index produced by the BLS measures housing prices solely by rents.

[19] One might consider constructing a price index with housing prices for each metro area measured by user costs of homeownership net of expected appreciation and then estimate the wage-price elasticity using this index. However, the benefits of this approach are limited. First, it is unclear how homeowners form expectations about appreciation and how homeowner user costs should be computed. Second, standard capital theory suggests that in order for the rental and homeowner housing markets to be in equilibrium, the user cost of rental housing should approximately equal the user cost of homeowner housing. Therefore, quality-adjusted gross rents may be the best available proxy for homeowner user costs.

[20] The ACCRA *Cost of Living Index* also reports average rents for an area, but for a number of reasons quality-adjusted rents from the ACS are likely preferable to rents from the ACCRA index.

[21] Gross rents include rents as well as basic utilities (water, electricity, and gas) and home heating fuels (wood, kerosene, oil, coal, etc.). These utilities are often included in rental payments for some renters, but not for others. Therefore, gross rents are more comparable across households because they include utilities and fuels for all renter households.

[22] If, however, there are unobserved aspects of housing quality that are correlated with wages in a city, the estimated wage-price elasticity may be upwardly biased.

[23] For these purposes, non-housing prices are computed as a weighted average of ACCRA sub-indices for groceries (0.13), transportation (0.25), healthcare (0.06), and miscellaneous goods and services (0.56). Note, that this excludes utilities in addition to housing because utilities are largely already included in gross rents.

[24] Note that this expenditure share for housing differs from official reports of the CES expenditure share for both "Housing" and "Shelter." The housing share based on gross rents used herein includes certain utilities but excludes others and also excludes expenditures for household operations, housekeeping, and household furnishings. The housing share of 0.29 also differs from the official CES tabulations in that homeowner housing expenditures are measured by implicit rents and not by out-of-pocket expenses such as mortgage interest.

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For the sake of comparison, I also compute a modified price index that measures housing prices by quality-adjusted housing values from the 2006 ACS computed in a manner similar to quality-adjusted gross rents. For this second modified price index, housing prices are given a weight of 0.23 because values do not include utilities and non-housing prices (now including utilities) are given a weight of 0.77.

Summary statistics for several price variables are reported in Table 1. As seen, the modified price index using gross rents is considerably less dispersed than both the baseline price index and the modified price index using housing values. Equivalently, housing values are more dispersed across cities than are gross rents. Non-housing prices are much less dispersed across cities than both rents and values, but there is still considerable variation in non-housing prices. Appendix Table A lists the 167 cities included in the sample and their value for the rent-based price index.

In addition to estimating Eq. (7), this paper is also interested in the relationship between wages and the prices of housing and non-housing goods and services. Therefore, I also divide the price index into housing prices, P_1, and non-housing prices, P_2, and include them in logarithmic form in the log wage equation separately:

$$\ln W_{ij} = X_{ij}\beta + \theta_1 \ln P_{1j} + \theta_2 \ln P_{2j} + \gamma Z_j + \varepsilon_{ij}. \quad (9)$$

Examining housing prices separately from non-housing prices is interesting for several reasons. For one, it allows us to test if the prediction of Eq. (4) holds for housing and non-housing prices separately. Additionally, a large literature in urban and regional economics following Roback (1982) ranks the quality of life across cities using implicit prices of amenities computed as the sum of compensating differentials in housing and labor markets, $\mathrm{d}P_1/\mathrm{d}Z - \mathrm{d}W/Z$. Few of these studies incorporate non-housing prices (Gabriel et al., 2003; Shapiro, 2006; Albouy, 2008b are recent exceptions). The justification for this exclusion is often that non-housing prices are relatively unimportant (Beeson and Eberts, 1989). The non-trivial variation in non-housing prices illustrated in Table 1 combined with the large budget share for non-housing consumption, however, suggests that non-housing prices may be quite important. The few papers that do incorporate non-housing prices often do so in a less than ideal way.[25] Separating housing and non-housing prices allows us to examine the importance of each in explaining interarea wage differentials.

Theory and previous empirical evidence predict that amenities also affect both wages and local prices. Therefore, my regressions also include a number of different amenities from several sources found to be important in previous literature.[26] A list of variables and data sources is included in Appendix Table B. Without including amenities, the estimated relationship between wages and prices could be biased.[27] Data for several natural amenities are obtained from the USDA Economic Research Service. These include the mean January temperature in degrees Fahrenheit, mean July temperature, mean hours of January sunlight, mean July relative humidity, the percent of land area covered by water, and five indicator variables for topography that range from very flat to mountainous. The flattest land surface is the omitted reference group. Mean annual inches of precipitation and snow are obtained from *Cities Ranked and Rated, 2nd Edition* (Sperling and Sandler, 2007). Maps were consulted to create indicator variables for whether a city is located on the coast of the Atlantic Ocean, Pacific Ocean or Gulf of Mexico. Data on violent crime and property crime per capita were obtained from the Census Bureau's USA Counties website. The mean commuting time in minutes for workers in a city was computed using the 2006 ACS microdata. Two measures of air pollution, ozone and particulate matter 2.5, were computed using the EPA AirData database.[28] The regressions also include eight census division dummies and six city population size dummies to account for residual differences in amenities.[29] The city size dummies should also help control for differences in unobserved worker ability across cities.[30] No specification of amenities is likely to fully capture differences in the quality of life across cities, but the hope is that the variables used in this paper do a reasonably good job of controlling for differences in the quality of life across cities.

Table 1
Summary statistics for price indices, 2006.

	Min.	Max.	St. Dev.
Baseline price index	84.0	157.9	12.0
Rent-based modified price index	84.1	141.8	9.2
Housing value-based modified price index	80.3	184.8	15.7
Quality-adjusted gross rents	66.4	184.4	20.0
Quality-adjusted housing values	46.9	395.0	52.9
Non-housing prices	86.7	124.4	5.7

Notes: Un-weighted mean is normalized to 100. Standard Deviation is un-weighted. Includes 167 cities.

5. Empirical results: the elasticity between wages and the general price level

This section presents results of the elasticity between wages and the general price level using the baseline price index, the price index modified using quality-adjusted gross rents, and the price index modified using quality-adjusted house values. All regressions include the full list of amenities, division dummies, city size dummies, and individual characteristics as explanatory variables. The results for these variables were generally as expected and are available upon request. I begin by estimating the regressions using Ordinary Least Squares and then proceed to instrument for prices to account for measurement error, which would bias the estimated coefficients toward zero. All of the price index coefficients in this section are statistically different from zero at the 1% level using cluster robust standard errors, but the more appropriate null hypothesis is whether or not they are different from unity.[31]

5.1. Ordinary Least Squares

I first estimate the wage-price elasticity, θ, using the baseline price index via OLS. This specification is comparable to that of DuMond et al. (1999), but my equation contains many more amenities, more recent data, and uses after-tax wages as the dependent variable.[32] As seen in

[25] For example, both Shapiro (2006) and Albouy (2008b) infer non-housing prices from housing prices by regressing non-housing prices on housing prices using the ACCRA *Cost of Living Index*. However, their approach ignores differences in non-housing prices across cities that are not correlated with housing prices. My own analysis suggests that regressing non-housing prices on division dummies, city size dummies, and amenities in addition to housing prices does a much better job of predicting non-housing prices than housing prices alone.

[26] Many of these are reported at the sub-metropolitan level and had to be aggregated to the CBSA/CSA level using populations as weights.

[27] For example, a pure consumption amenity is likely to drive up housing prices and drive down wages, which would bias the wage-price elasticity toward zero.

[28] Pollution values were unavailable for several small cities and were imputed based on average values by Census division and city size. Particulate matter was imputed in this manner for 16 cities, and ozone was imputed for 23 cities. I tested the potential effect of this imputation by estimating the regressions without pollution variables and estimating the regressions with pollution variables but only for cities that had unimputed pollution levels. The main results of this paper do not appear to be affected by the imputation of pollution values for these small cities.

[29] The seven city size categories are: 0–199,999; 200,000–299,999; 300,000–499,999; 500,000–999,999; 1,000,000–1,999,999; 2,000,000–4,999,999; and 5,000,000+.

[30] Glaeser and Maré (2001), Yankow (2006), and Krupka (2007) all find that the nominal city size wage premium falls after controlling for individual fixed effects using panel data on workers, suggesting that large cities attract more able workers.

[31] Unless otherwise noted, all standard errors in this paper are clustered by city.

[32] A more subtle difference is that DuMond et al. (1999) include workers with imputed earnings, which likely biases their estimates toward zero.

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Table 2
OLS results for three price indices.

	1	2	3
Log baseline index	0.314[a] (0.048)		
Log rent-based modified index		0.760[a] (0.078)	
Log value-based modified index			0.416[a] (0.049)
R^2	0.494	0.495	0.494

Notes: Dependent variable is the log of hourly wages net of federal income taxes computed from the 2006 CPS-ORG files. Standard errors in parentheses are robust, clustered by CSA/CBSA. Regressions contain observations on 71,705 workers in 167 cities. Regressions also include 8 Census division dummies, 6 city size dummies, January temperature, July temperature, January sun, July humidity, the % of land area covered by water, 4 indicators for topography, 3 indicators for coastal location, precipitation, snow, violent crime, property crime, ozone, particulate matter (2.5), mean commute time, 11 education dummies, a quartic specification for experience, dummy variables for whether a worker is female, Black, Asian, Hispanic, Other, married, employed part-time, enrolled part-time in school, a member of a union, a naturalized citizen, or a non-citizen, 9 occupation dummies, 11 industry dummies, 3 dummies for government employment, and 11 month in sample dummies. The baseline index refers to the price index constructed solely using ACCRA data. The two modified indices combine housing prices from the Census with non-housing prices from ACCRA. See text for further details.

[a] Significantly different from unity at the 1% level.

the first column of Table 2, this specification yields an estimate of θ of 0.314, and the coefficient is statistically different from one at the 1% level. According to this estimate, a one percent increase in the general price level in a city is associated with a 0.31% increase in after-tax wages. This is also considerably lower than the previous estimate of 0.46 by DuMond et al. (1999). This may suggest that the sharp increase in housing values in recent years causes the ACCRA index to be a worse measure of the cost of living in 2006 than it was between 1985 and 1995, the time period considered by DuMond et al. (1999).

The baseline index, however, may do a poor job of measuring differences in prices across cities in part because it measures housing prices primarily according to house values instead of rents. Therefore, the rent-based modified price index, which measures housing prices solely by gross rents from the ACS, may be more appropriate. Using the rent-based price index, OLS yields an estimated wage-price elasticity of 0.760, much higher than for the baseline price index. This is an important result. It appears that the wage-price elasticity using the baseline price index is biased toward zero in part because of how housing prices are measured. However, the estimate for the rent-based index is still significantly less than one.

I also estimate θ using the housing value-based modified price index. Using OLS, the estimated coefficient is 0.416 and is significantly less than one. Interestingly, the coefficient for the value-based modified index is greater than that for the baseline index. This suggests that measuring housing prices by values may not be the only source of measurement error in the baseline index.[33]

5.2. Instrumental variables

Even after measuring housing prices by quality-adjusted gross rents from the ACS, the rent-based price index may still be measured with considerable error. Gross rents in the ACS are likely subject to some degree of sampling error and non-housing prices measured in the ACCRA *Cost of Living Index* may be subject to a number of sources of measurement error. Random measurement error will bias the coefficient on the log of the price index toward zero, and including variables that are highly correlated with the price index such as amenities, division dummies, and city size dummies, may exacerbate measurement error bias. I next use instrumental variables to account for measurement error in the rent-based price index. I use as instruments lagged values of log gross rents and log non-housing prices. If measurement error is random, then instrumenting for the price index using the previous year's components should produce consistent estimates of θ. If measurement error in the price index is serially correlated, however, instrumenting using lagged prices will not produce consistent coefficient estimates.

Table 3 presents 2SLS results for the rent-based modified index. The instruments used are highly significant in the first-stage regressions reported in the lower half of the table. I first instrument for the log of the rent-based price index using quality-adjusted log gross rents from the previous year. As reported in the first column of Table 3, instrumenting in this manner yields a coefficient estimate of 0.994 that is nearly identical to one. Therefore, instrumenting for the rent-based price index using rents for the previous year provides empirical support for the full compensation hypothesis. I next instrument for the rent-based modified index using non-housing prices for the previous year. The 2SLS coefficient estimate in this case, 0.603, is considerably lower than that found using OLS. Finally, when we use both gross rents and non-housing prices as instruments for the rent-based price index, we get a coefficient estimate of 0.830 that is statistically different from unity at the 10% level.

Non-housing prices are constructed from the ACCRA *Cost of Living Index* and are likely subject to considerable measurement error, some of which is likely persistent within cities over time. If measurement error in non-housing prices is serially correlated, then instrumenting for the general price level using non-housing prices will not yield consistent estimates of θ. The divergence between the estimates in the first and second columns of Table 3 suggests that this is indeed the case. Quality-adjusted gross rents are estimated from the ACS PUMS and may also be subject to some measurement error such as due to sampling. However, the measurement error in log gross rents is much more likely to be classical in nature. If the measurement error in the lag of log gross rents is purely random and uncorrelated with measurement error in the rent-based price index, then the 2SLS estimates in the first column of Table 3 are consistent. This seems quite plausible. If log gross rents are a valid instrument, over-identification in the specification of the third column allows us to examine the validity of non-housing prices as an instrument. Doing so, we get a Hansen J Statistic of 11.297, which allows us to reject non-housing prices as a valid instrument at the 1% level. Thus the coefficient in the first column of Table 3 is the preferred estimate of the elasticity between wages and the general price level.

In results not shown, I also estimate the wage-price elasticity for the baseline index and the value-based modified price index using 2SLS with lagged price index components as instruments. The 2SLS

Table 3
2SLS results for the rent-based modified index.

	1	2	3
Second-stage results			
Log rent-based modified index, 2006	0.994 (0.106)	0.603[b] (0.108)	0.830[a] (0.091)
R^2	0.494	0.495	0.495
First-stage results			
Log gross rents, 2005	0.377*** (0.024)		0.290*** (0.017)
Log non-housing index, 2005		0.878*** (0.077)	0.594*** (0.047)
Partial R^2 of excluded instruments	0.657	0.526	0.859

Notes: Regression in column 1 contains observations on 71,705 workers in 167 cities, while regressions in columns 2 and 3 contain observations on 69,743 workers in 157 cities. The dependent variable and additional regressors are the same as in Table 2. Standard errors in parentheses are robust, clustered by CSA/CBSA.

[a]Significantly different from unity at the 10% level. [b]Significantly different from unity at the 1% level. ***Significantly different from zero at the 1% level in the first-stage regressions.

[33] It may also be the case that housing values are measured with greater error than rents and this leads to greater measurement error bias for the value-based modified index than the rent-based index. Bucks and Pence (2006), however, report that homeowner reported housing values are fairly accurate.

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Table 4
2SLS results for the rent-based modified index under alternative specifications.

	Coefficient	Standard error
(1) Preferred specification	0.994	0.106
(2) Pre-tax wages	1.062	0.114
(3) Including state fixed effects	0.949	0.110
(4) Renters only	1.037	0.128
(5) Homeowners only	1.011	0.122

Notes: Results are from 2SLS regressions for the log of the rent-based price index using log gross rents for 2005 as an instrument. Regressions contain observations on 71,705 workers in 167 cities. The dependent variable and additional regressors are the same as in Table 2. Standard errors are robust, clustered by CSA/CBSA. None of the coefficients is statistically different from unity at usual levels of significance.

elasticities for the baseline and value-based indices are slightly greater than the corresponding OLS estimates, but never greater than 0.5.

A recap of the results in this section is warranted. When housing prices are measured by homeowner values, the estimated elasticity between wages and the general price level is never more than 0.5, even when I use instrumental variables to account for measurement error. When housing prices are measured by rents, though, the estimated elasticity between wages and the general price level increases considerably. Using OLS the estimated wage-price elasticity is 0.76, but instrumenting for the log of the rent-based price index using the log of quality-adjusted gross rents for the previous year, the wage-price elasticity is equal to one for all practical purposes. This result supports the full compensation hypothesis and has important implications for researchers estimating the implicit prices of amenities. In the next section, I examine the sensitivity of θ to alternative specifications.

6. The elasticity between wages and the general price level for alternative specifications

In this section, I briefly examine 2SLS wage-price elasticity estimates using the rent-based price index under some alternative specifications. The results are presented in Table 4. In all specifications the rent-based price index is instrumented for using my preferred instrument, the log of quality-adjusted gross rents in 2005. The first row of Table 4 reproduces estimates for the preferred specification from the first column of Table 3.

6.1. Pre-tax wages

In the second row of Table 4, I estimate θ via 2SLS using pre-tax wages as the dependent variable. As pointed out by Henderson (1982) and Albouy (2008a,b), the progressivity of the federal income tax causes workers in cities with high nominal wages to pay a higher percentage of their income in federal income taxes than workers in cities with lower nominal wages. For the utility of homogenous workers to be constant across areas, pre-tax wages should be more dispersed across areas than after-tax wages. In other words, workers in high wage areas must be compensated for the higher federal income taxes they pay in addition to the compensation they require for the higher cost of living or worse bundle of amenities. As such, the estimated wage-price elasticity should be higher using pre-tax wages than using after-tax wages. The results in row 2 suggest that this is indeed the case. The estimate of θ increases to 1.062, but is not statistically different from unity. I maintain, however, that it is after-tax wages that should equalize across areas controlling for prices, amenities, and individual characteristics, so measuring wages net of federal income tax provides a better test of the theory than measuring wages before federal income tax.

6.2. State fixed effects

If wages should be measured net of federal income tax, we might also consider adjusting wages for state and local income taxes. Income taxes, however, are only part of the story at the states and local level. To adjust wages for state and local income taxes, we would also need to incorporate information on other state and local taxes and state and local public spending. To avoid the many complexities involved with adjusting wages for state and local taxes and expenditures, I adopt a different approach by examining the robustness of my results to including state fixed effects. If ignoring state taxes and expenditures is biasing the previous results, then we would expect that including state fixed effects would produce a very different estimate of θ than the case in which we include census division fixed effects. As seen in row 3, including state fixed effects reduces the coefficient estimate to 0.949, but it is not statistically different from one. Therefore, the basic findings of this paper are robust to including state fixed effects. My preferred specification, however, is to use census division dummies and not state fixed effects because several states contain only one city in the sample. Including state fixed effects means that θ is only estimated based on states that have more than one city in the sample.

6.3. Renters vs. homeowners

One might also wonder if using a rent-based price index yields different estimates of θ for renters and homeowners. Rows 4 and 5 of Table 4 estimate θ separately for renters and homeowners. The coefficient estimate for renters is 1.037, and the estimate for homeowners is 1.011. Therefore, the coefficient estimates for renters and homeowners separately are slightly higher than the pooled estimate, but neither estimate is statistically different from unity. The coefficient estimates for renters and homeowners are also not statistically different from each other. It appears that differences in prices across cities affect the wages of renters and homeowners roughly the same.

7. Empirical results: the elasticity between wages and housing and non-housing prices

I next separate the price index into housing and non-housing prices and include them in the wage equation separately. The model predicts that under certain conditions the wage-price elasticity for a good should be approximately equal to the expenditure share for the good. I wish to explore the validity of this hypothesis for both housing and non-housing prices. Therefore, based on expenditure shares computed from the 2005 CES, the expected coefficient for housing is about 0.29 and the expected coefficient for non-housing is roughly 0.71.

Table 5
Separating housing prices (rents) and non-housing prices.

	1	2	3
	OLS	2SLS	2SLS
Full adjustment for	N/A	Non-housing prices	Housing prices
Log gross rents	0.337 (0.038)	0.297 (0.042)	
Log non-housing price index	0.231[a] (0.106)		0.754 (0.289)
R^2	0.495	0.483	0.482
First-stage results			
Log gross rents, 2005		0.934*** (0.032)	0.138*** (0.029)
Partial R^2 of excluded instruments		0.863	0.141

Notes: The dependent variable in column 1 is the log of after-tax hourly wages. In column 2 wages are fully adjusted for non-housing prices, i.e. the coefficient on log non-housing prices is constrained to equal 0.71. In column 3 wages are fully adjusted for housing prices measured by gross rents, i.e. the coefficient on log gross rents is constrained to equal 0.29. All regressions contain observations on 71,705 workers in 167 cities. The additional regressors are the same as in Table 2. Standard errors in parentheses are robust, clustered by CSA/CBSA.
[a]Significantly different from the budget share (0.29 for housing and 0.71 for non-housing) at the 1% level. ***Significantly different from zero at the 1% level in the first-stage regressions.

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I first estimate the log wage equation with log gross rents and log non-housing prices included simultaneously via OLS. As discussed above, measurement error in prices may bias coefficients toward zero. Alternatively, if non-housing prices are measured with considerable error, while housing prices are measured with relatively little error, the coefficient on housing prices could be biased upward from picking up some of the effect of non-housing prices. This is especially problematic given the very high correlation between log gross rents and log non-housing prices; the raw correlation coefficient between the two is 0.718. The results in column 1 of Table 5 suggest that log gross rents may indeed be picking up some of the effect of log non-housing prices. The coefficient on log gross rents is 0.337, and is statistically different from zero at the 1% level but not statistically different from the budget share of 0.29. The coefficient on log non-housing prices is 0.231, and is statistically different from zero at the 5% level and statistically different from the budget share of 0.71 at the 1% level.

Ideally, we would like to simultaneously instrument for housing and non-housing prices to account for measurement error in both. One possibility would be to use lagged values of both as instruments. However, because measurement error in non-housing prices is likely to be serially correlated, instrumenting for non-housing prices using its lagged value will not yield consistent estimates. Instead, I explore estimating the log housing price and log non-housing price coefficients separately while constraining the other to equal its budget share and instrumenting using log gross rents for the previous year. This is a hybrid between full adjustment and partial adjustment for prices used by previous researchers. Constraining one of the coefficients to be different from its true value, however, will likely bias the other in the opposite direction. First-stage results at the bottom of Table 5 confirm that the log of gross rents from the previous year is a significant predictor of both log gross rents and log non-housing prices.

In column 2 of Table 5, wages are fully adjusted for non-housing prices by constraining the coefficient on log non-housing prices to equal 0.71, i.e., I estimate:

$$\ln W_{ij} - .71^{*} \ln P_{2j} = X_{ij}\beta + \theta_1 \ln P_{1j} + \gamma Z_j + \varepsilon_{ij}. \tag{10}$$

The coefficient on log gross rents is estimated by 2SLS using log gross rents for the previous year as an instrument. As seen, the coefficient on log gross rents falls to 0.297 and is not statistically different from 0.29. In other words, when we fully adjust wages for non-housing prices, the elasticity between wages and housing prices (measured by gross rents) is nearly identical to housing's budget share consistent with the prediction of the model.

Table 6
Separating housing prices (values) and non-housing prices.

	1	2	3
	OLS	2SLS	2SLS
Full adjustment for	NA	Non-housing prices	Housing prices
Log housing values	0.143[a] (0.024)	0.091[a] (0.021)	
Log non-housing price index	0.165[a] (0.132)		−0.641[a] (0.226)
R^2	0.494	0.482	0.484
First-stage results			
Log housing values, 2005		0.992*** (0.024)	0.097*** (0.015)
Partial R^2 of excluded instruments		0.949	0.281

Notes: The dependent variable in column 1 is the log of after-tax hourly wages. In column 2 wages are fully adjusted for non-housing prices, i.e. the coefficient on log non-housing prices is constrained to equal 0.77. In column 3 wages are fully adjusted for housing prices measured by housing values, i.e. the coefficient on log housing values is constrained to equal 0.23. All regressions contain observations on 71,705 workers in 167 cities. The additional regressors are the same as in Table 2. Standard errors in parentheses are robust, clustered by CSA/CBSA.
[a]Significantly different from the budget share (0.23 for housing and 0.77 for non-housing) at the 1% level. ***Significantly different from zero at the 1% level in the first-stage regressions.

Table 7
Amenity values by census division.

Wages fully adjusted using	Rent-based modified price index	Baseline price index	Gross rents only
Middle Atlantic	−0.017 (0.019)	−0.044 (0.042)	−0.020 (0.020)
East North Central	0.075*** (0.018)	0.128*** (0.019)	−0.019 (0.012)
West North Central	0.070*** (0.017)	0.160*** (0.022)	−0.038** (0.018)
South Atlantic	0.041** (0.020)	0.112*** (0.030)	−0.033 (0.020)
East South Central	0.059*** (0.017)	0.132*** (0.022)	−0.045*** (0.014)
West South Central	0.065*** (0.020)	0.203*** (0.032)	−0.054*** (0.016)
Mountain	0.072*** (0.020)	0.159*** (0.021)	−0.018 (0.015)
Pacific	−0.004 (0.022)	−0.036 (0.041)	−0.003 (0.020)

Notes: Regressions contain detailed individual characteristics as in Table 2, but no city-level variables other than Census division dummies. New England is the reference group. Standard errors in parentheses are robust, clustered by CSA/CBSA.
**Significantly different from zero at the 5% level. ***Significantly different from zero at the 1% level.

In column 3 of Table 5, wages are fully adjusted for housing prices by constraining the coefficient on log gross rents to equal 0.29:

$$\ln W_{ij} - .29^{*} \ln P_{1j} = X_{ij}\beta + \theta_2 \ln P_{2j} + \gamma Z_j + \varepsilon_{ij}. \tag{11}$$

The coefficient for log non-housing prices is estimated by 2SLS using log gross rents for the previous year as an instrument. Obviously, if the true value of θ_1 is greater than 0.29, the estimate for θ_2 will be upwardly biased. That said, the coefficient on log non-housing prices is 0.754 and is not statistically different from 0.71. The 2SLS results in Table 5, therefore, suggest that the prediction of the model that the wage-price elasticity for a good is equal to its budget share holds for housing and non-housing prices separately.

In Table 6, I reestimate the regressions in Table 5 measuring housing prices by quality-adjusted housing values.[34] The coefficients on log housing values and log non-housing prices are always significantly less than their budget shares. In fact, when I fully adjust wages for housing prices measured by housing values in column 3 of Table 6, the log of the non-housing price index has a significantly negative coefficient. This reinforces results in the previous section suggesting that housing values are an inappropriate measure of housing prices.

8. Implications for estimating implicit prices of amenities

The empirical results in this paper have important implications for researchers interested in estimating the implicit prices of amenities or ranking the quality of life across cities. The relationship between wages and prices is consistent with the full compensation hypothesis when we measure housing prices by rents and use lagged rents as an instrument for the general price level. When we measure housing prices by values, however, the relationship between wages and prices is highly inconsistent with the full compensation hypothesis, even using instrumental variables. This suggests that using housing values along with wages to infer implicit prices of amenities is likely to produce biased estimates. To illustrate, I estimate Census division amenity values by regressing log after-tax wages fully adjusted by both the rent-based modified price index and the baseline price index on eight Census division dummy variables. These regressions contain individual worker characteristics but no city level controls other than Census division indicators. The results are presented in columns 1 and 2 of Table 7.[35] The implicit price of a division's amenities is measured as the negative of its division dummy coefficient for fully-adjusted

[34] The expected shares for housing and non-housing now change to 0.23 and 0.77 because housing values do not include utilities and non-housing prices now do.

[35] I also regressed the log of the rent-based modified index on log gross rents, amenities, region dummies, and city size dummies to obtain predicted values that "net out" potential measurement error. Division dummies estimated for wages fully adjusted using the predicted values of the rent-based modified priced index were nearly identical to those in column 1 using the actual values.

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wages. In other words, a low division coefficient indicates a high value of amenities.

If the true wage-price elasticity is equal to one and the rent-based modified index measures the general price level across cities without systematic error (but potentially random error), then the estimates of amenity values by division in column 1 are consistently estimated.[36] The estimates in column 1 suggest that the Middle Atlantic, Pacific, and New England (the omitted category) divisions have the most highly valued bundles of amenities. A coefficient of 0.065 for the West South Central division suggests that a marginal worker will require a 6.5% higher "real wage" to live in the West South Central division than in New England to compensate for the worse bundle of amenities.

Fully adjusting wages using the baseline index, however, may upwardly bias estimates of amenity values in areas with high values of the index and downwardly bias estimates of amenity values in areas with low values of the index. This result follows because housing values are more dispersed than rents across cities, but rents measure the true user cost of housing. The rank ordering of division dummies in column 2 is similar to that in column 1, but the estimated coefficients are much larger. According to wages fully adjusted using the baseline index, a marginal worker will require a more than 20% higher "real wage" to live in the West South Central than in New England. However, because the baseline index measures housing prices primarily by housing values, the estimated amenity prices in column 2 are biased.

This paper also has implications for researchers who neglect to include non-housing prices in measuring the implicit price of amenities. Column 3 of Table 7 reports the results of division dummies for log wages fully adjusted for gross rents (assuming a budget share of 0.29) but not non-housing prices. The ranking of the coefficients is nearly the opposite of that in column 1. The West South Central is now the most amenable and the New England and Pacific divisions are now the worst. These results confirm that ignoring non-housing prices downwardly biases amenity values for areas with high non-housing prices and upwardly biases amenity prices for areas with low non-housing prices.[37]

9. Conclusion

Differences in wages across areas can be partially explained by differences in prices and amenities. For a given price level, workers are willing to accept lower wages to live and work in more amenable locations. Controlling for amenities, wages must be higher in high price areas in order for workers to achieve equal utility across locations. This paper presents a simple model that predicts that the elasticity of the wage with respect to the price of a good is proportional to the share of wage income spent on the good. The model also suggests that if workers' consumption equals their wage income, then the elasticity between wages and the general price level should equal one. However, to the extent that the assumptions of the model do not hold, the actual relationship between wages and prices may differ from that predicted by the model.

Measuring housing prices by rents, I find that the elasticity between wages and the general price level is nearly identical to one after instrumenting for the general price level using rents for the previous year. I also present evidence that the wage-price elasticities for housing and non-housing prices are equal to their budget shares when housing prices are measured by rents. These results provide empirical support for the full compensation hypothesis.

Importantly, though, when housing prices are measured by housing values, the elasticity between wages and the general price level is less than 0.5. The findings in this paper have important implications for estimating the implicit prices of amenities. Measuring housing prices by values instead of rents will bias estimates and cause cities with high housing values to have the relative value of their amenities overstated.

Appendix A

Appendix Table A
Rent-based price index by city, 2006.

City	Index	City	Index
Albany-Schenectady-Amsterdam, NY CSA	108.3	Davenport-Moline-Rock Island, IA-IL CBSA	94.7
Albuquerque, NM CBSA	101.1	Dayton-Springfield-Greenville, OH CSA	96.2
Amarillo, TX CBSA	91.9	Decatur, IL CBSA	90.8
Anniston-Oxford, AL CBSA	92.7	Denver-Aurora-Boulder, CO CSA	105.4
Appleton-Oshkosh-Neenah, WI CSA	99.6	Des Moines-Newton-Pella, IA CSA	95.0
Asheville-Brevard, NC CSA	97.1	Detroit-Warren-Flint, MI CSA	101.6
Atlanta-Sandy Springs-Gainesville, GA-AL CSA	105.8	Dover, DE CBSA	103.7
Augusta-Richmond County, GA-SC CBSA	94.5	Eau Claire-Menomonie, WI CSA	92.7
Austin-Round Rock, TX CBSA	107.0	El Paso, TX CBSA	91.9
Bakersfield, CA CBSA	105.2	Erie, PA CBSA	95.4
Bangor, ME CBSA	100.0	Eugene-Springfield, OR CBSA	105.2
Baton Rouge-Pierre Part, LA CSA	96.8	Evansville, IN-KY CBSA	92.0
Beaumont-Port Arthur, TX CBSA	94.2	Fargo-Wahpeton, ND-MN CSA	91.8
Bellingham, WA CBSA	104.7	Farmington, NM CBSA	94.0
Bend-Prineville, OR CSA	110.2	Fayetteville, NC CBSA	100.8
Birmingham-Hoover-Cullman, AL CSA	98.7	Fayetteville-Springdale-Rogers, AR-MO CBSA	93.0
Bloomington, IN CBSA	98.9	Florence-Muscle Shoals, AL CBSA	87.3
Bloomington-Normal, IL CBSA	99.1	Fort Collins-Loveland, CO CBSA	105.6
Boise City-Nampa, ID CBSA	100.6	Fort Smith, AR-OK CBSA	87.9
Boston-Worcester-Manchester, MA-RI-NH CSA	123.2	Fort Walton Beach-Crestview-Destin, FL CBSA	104.6
Bowling Green, KY CBSA	92.2	Fort Wayne-Huntington-Auburn, IN CSA	93.3
Brownsville-Harlingen-Raymondville, TX CSA	87.8	Fresno-Madera, CA CSA	107.9
Buffalo-Niagara-Cattaraugus, NY CSA	100.3	Gainesville, FL CBSA	101.4
Burlington-South Burlington, VT CBSA	115.9	Grand Rapids-Muskegon-Holland, MI CSA	99.1
Cape Coral-Fort Myers, FL CBSA	112.7	Green Bay, WI CBSA	97.4
Cedar Rapids, IA CBSA	94.4	Greensboro-Winston-Salem, NC CSA	94.5
Champaign-Urbana, IL CBSA	99.4	Greenville-Spartanburg-Anderson, SC CSA	92.8
Charleston, WV CBSA	88.5	Gulfport-Biloxi-Pascagoula, MS CSA	99.1
Charleston-North Charleston, SC CBSA	103.4	Hagerstown-Martinsburg, MD-WV CBSA	94.0
Charlotte-Gastonia-Salisbury, NC-SC CSA	98.2	Harrisburg-Carlisle-Lebanon, PA CSA	100.7
Chattanooga-Cleveland-Athens, TN-GA CSA	90.6	Harrisonburg, VA CBSA	97.6
Chicago-Naperville-Michigan City, IL-IN-WI CSA	112.5	Hartford-West Hartford-Willimantic, CT CSA	112.8
Cincinnati-Middletown-Wilmington, OH-KY-IN CSA	95.6	Hickory-Lenoir-Morganton, NC CBSA	92.8
Cleveland-Akron-Elyria, OH CSA	98.9	Houston-Baytown-Huntsville, TX CSA	102.0
Colorado Springs, CO CBSA	102.7	Huntsville-Decatur, AL CSA	94.2

[35] I also regressed the log of the rent-based modified index on log gross rents, amenities, region dummies, and city size dummies to obtain predicted values that "net out" potential measurement error. Division dummies estimated for wages fully adjusted using the predicted values of the rent-based modified priced index were nearly identical to those in column 1 using the actual values.

[36] The previous two sections argue that there is systematic measurement error within cities over time in non-housing prices. This measurement error, however, can still be unsystematic across cities for a given time period.

[37] Interestingly, though, the biases from measuring housing prices by housing values and ignoring non-housing prices are in opposite direction. As a result, measuring housing prices by values and ignoring non-housing prices produces amenity estimates generally between those in column 2 and column 3.

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Appendix Table A (*continued*)

City	Index	City	Index
Columbia, MO CBSA	95.4	Indianapolis-Anderson-Columbus, IN CSA	97.9
Columbia-Newberry, SC CSA	97.4	Jacksonville, FL CBSA	92.9
Columbus-Auburn-Opelika, GA-AL CSA	98.1	Jacksonville, NC CBSA	103.2
Columbus-Marion-Chillicothe, OH CSA	101.1	Jackson-Yazoo City, MS CSA	96.4
Corpus Christi-Kingsville, TX CSA	97.2	Janesville, WI CBSA	97.4
Dallas-Fort Worth, TX CSA	104.4	Johnson City-Kingsport-Bristol, TN-VA CSA	86.0
Johnstown, PA CBSA	88.6	Port St. Lucie-Sebastian-Vero Beach, FL CSA	108.5
Joplin, MO CBSA	84.1	Portland-Lewiston-South Portland, ME CSA	106.2
Kalamazoo-Portage, MI CBSA	95.8	Portland-Vancouver-Beaverton, OR-WA CBSA	111.4
Kansas City-Overland Park, MO-KS CSA	99.3	Prescott, AZ CBSA	106.1
Killeen-Temple-Fort Hood, TX CBSA	95.8	Pueblo, CO CBSA	90.6
Knoxville-Sevierville-La Follette, TN CSA	88.7	Raleigh-Durham-Cary, NC CSA	101.8
Lafayette-Acadiana, LA CSA	93.2	Reno-Sparks-Fernley, NV CSA	111.0
Lake Charles-Jennings, LA CSA	92.2	Richmond, VA CBSA	105.1
Lancaster, PA CBSA	105.8	Roanoke, VA CBSA	94.7
Laredo, TX CBSA	86.4	Rochester-Batavia-Seneca Falls, NY CSA	105.7
Las Cruces, NM CBSA	94.6	Sacramento-Arden-Truckee, CA-NV CSA	116.1
Las Vegas-Paradise-Pahrump, NV CSA	109.2	Salt Lake City-Ogden-Clearfield, UT CSA	103.1
Lawrence, KS CBSA	98.4	San Antonio, TX CBSA	97.5
Lawton, OK CBSA	88.8	San Diego-Carlsbad-San Marcos, CA CBSA	133.8
Lexington-Fayette-Frankfort-Richmond, KY CSA	92.7	San Jose-San Francisco-Oakland, CA CSA	141.8
Little Rock-North Little Rock-Pine Bluff, AR CSA	93.0	Sarasota-Bradenton-Punta Gorda, FL CSA	110.6
Longview-Marshall, TX CSA	91.1	Savannah-Hinesville-Fort Stewart, GA CSA	103.0
Los Angeles-Long Beach-Riverside, CA CSA	128.4	Seattle-Tacoma-Olympia, WA CSA	114.9
Louisville-Elizabethtown-Scottsburg, KY-IN CSA	97.5	Shreveport-Bossier City-Minden, LA CSA	92.7
Lubbock-Levelland, TX CSA	95.1	Sioux Falls, SD CBSA	94.3
Macon-Warner Robins-Fort Valley, GA CSA	94.3	South Bend-Elkhart-Mishawaka, IN-MI CSA	94.2
McAllen-Edinburg-Mission, TX CBSA	87.1	Spokane, WA CBSA	102.3
Memphis, TN-MS-AR CBSA	98.7	Springfield, IL CBSA	95.3
Miami-Fort Lauderdale-Miami Beach, FL CBSA	118.1	Springfield, MO CBSA	93.1
Midland-Odessa, TX CSA	94.1	St. Louis-St. Charles-Farmington, MO-IL CSA	97.9
Milwaukee-Racine-Waukesha, WI CSA	102.7	Stockton, CA CBSA	113.5
Minneapolis-St. Paul-St. Cloud, MN-WI CSA	107.4	Syracuse-Auburn, NY CSA	101.8
Mobile-Daphne-Fairhope, AL CSA	95.4	Tampa-St. Petersburg-Clearwater, FL CBSA	106.4
Montgomery-Alexander City, AL CSA	94.7	Toledo-Fremont, OH CSA	95.3
Myrtle Beach-Conway-Georgetown, SC CSA	97.2	Topeka, KS CBSA	93.6
Nashville-Murfreesboro-Columbia, TN CSA	98.9	Tucson, AZ CBSA	103.0
New Orleans-Metairie-Bogalusa, LA CSA	105.6	Tulsa-Bartlesville, OK CSA	95.0
New York-Newark-Bridgeport, NY-NJ-CT-PA CSA	132.0	Tuscaloosa, AL CBSA	97.8
Norwich-New London, CT CBSA	116.1	Valdosta, GA CBSA	92.7
Oklahoma City-Shawnee, OK CSA	95.4	VA Beach-Norfolk-Newport News, VA CBSA	104.4
Omaha-Council Bluffs-Fremont, NE-IA CSA	95.5	Waco, TX CBSA	92.3
Orlando-Deltona-Daytona Beach CSA	111.0	Washington-Baltimore, DC-MD-VA-WV CSA	121.5
Panama City-Lynn Haven, FL CBSA	106.3	Waterloo-Cedar Falls, IA CBSA	90.9

Appendix Table A (*continued*)

City	Index	City	Index
Pensacola-Ferry Pass-Brent, FL CBSA	100.0	Wausau-Merrill, WI CSA	94.7
Peoria-Canton, IL CSA	96.1	Wichita-Winfield, KS CSA	94.7
Philadelphia-Camden-Vineland, PA-NJ-DE-MD CSA	117.0	York-Hanover-Gettysburg, PA CSA	97.9
Phoenix-Mesa-Scottsdale, AZ CBSA	106.8	Youngstown-Warren, OH-PA CSA	89.9
Pittsburgh-New Castle, PA CSA	94.3		

Appendix B

Appendix Table B
Variables and data sources.

Variable	Data source
Log wage	Current Population Survey
Worker characteristics	Current Population Survey
Baseline price index	ACCRA
Rent-based modified price index	American Community Survey & ACCRA
Housing value-based modified price index	American Community Survey & ACCRA
Quality-adjusted gross rents	American Community Survey
Quality-adjusted housing values	American Community Survey
Non-housing prices	ACCRA
Gulf Coast	Consulted Map
Atlantic Coast	Consulted Map
Pacific Coast	Consulted Map
January temperature	ERS Natural Amenities Scale
July temperature	ERS Natural Amenities Scale
January sun	ERS Natural Amenities Scale
July humidity	ERS Natural Amenities Scale
% Water area	ERS Natural Amenities Scale
Topography 2	ERS Natural Amenities Scale
Topography 3	ERS Natural Amenities Scale
Topography 4	ERS Natural Amenities Scale
Topography 5	ERS Natural Amenities Scale
Precipitation	Cities Ranked and Rated
Snow	Cities Ranked and Rated
Violent crime	USA Counties Website
Property crime	USA Counties Website
Mean commute time	American Community Survey
Ozone	EPA AirData Database
Particulate matter (2.5)	EPA AirData Database
Census division indicators	Assigned According to Census Geography
City size indicators	Population Estimates from Census Bureau

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