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Car Ownership, Employment, and Earnings

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Abstract

We assess whether the positive relationship between car ownership and employment outcomes reflect a causal effect of auto ownership. We match state-level data on car insurance premiums and gas taxes to a microdata sample containing information on car ownership and employment outcomes. In OLS regressions that control for observable covariates, we find large differences in employment rates, weekly hours worked, and hourly earnings between those with and without cars. Instrumenting for car ownership with insurance and gas taxes yields estimates of the employment and hours effects that are quite close to the OLS estimates. Concerning wages, instrumenting eliminates the positive impact of auto ownership.

JEL Codes: E24, J22, J61, R41

Suggested Running Head: Car Ownership and Employment Outcomes

1. Introduction

Regardless of where one resides within a metropolitan area, having access to a car is likely to afford tangible advantages in locating and maintaining employment. For inner-city residents, the continuing spatial decentralization of total employment within metropolitan areas (Kasarda [8], HUD [17]) and, in particular, of low-skilled employment opportunities (Stoll, Holzer, & Ihlanfeldt [15]), may render car ownership a virtual necessity for accessing distant job centers. Even for workers residing in suburban communities, the spatially diffuse patterns of suburban economies may be more amenable to commuting by private auto rather than public transit. In light of these considerations, several researchers have argued for public policy that encourages car-ownership among the low and moderately-skilled (Ong [10], Ong & Blumenberg [11], O'Regan & Quigley [12]). In particular, Ong & Blumenberg [11] suggest that such policies should be an integral component of programs intended to move welfare recipients into sustainable employment.

The evidence offered in support of car-ownership policies consists of empirical studies demonstrating that individuals who own cars are more likely to be employed and, conditional on being employed, earn more than individuals without cars (Holzer et. al. [7], Ong [10]). While these findings are consistent with causal effects of car ownership, there are also several alternative explanations. For example, causation may run in the opposite direction; namely, those with jobs can afford cars. Alternatively, car-ownership may be determined in part by unobserved factors that also affect employability -- e.g., skills, or motivation. Given these highly plausible alternatives, the ability of the existing research to inform policy makers is severely limited.

In this paper, we assess whether the positive effects of car ownership on employment outcomes observed in past research are causal. We match state data on average car insurance premiums and per-gallon gas taxes to a nationally representative microdata sample containing information on car ownership and employment outcomes. State-level measures of insurance and fuel costs are unlikely to be correlated with unobserved skills or motivation yet are strongly correlated with car ownership rates. Hence, these variables provide the necessary exogenous variation in car-ownership rates for an analysis of the effect of having access to a car on labor market outcomes.

We use data from the 1992 and 1993 Survey of Income and Program Participation (SIPP). In OLS regressions that control for observable demographic and human capital variables, we find large differences in employment rates, weekly hours worked, and hourly earnings between those with and without cars. Instrumenting car ownership on insurance and gas tax costs yields estimates of the employment and hours effects of car ownership that are quite close to the OLS estimates. Concerning wages, the IV models yield negative effects of car ownership on wages. Despite these wage results, the employment and hours findings suggest important causal effects of car ownership on employment outcomes.

2. Why Should Cars Matter? Theory and a Brief Review of Previous Research

Several arguments suggest that access to reliable private transportation positively affects employment prospects. To start, commute times are lower using private transportation (Holzer et. al. [7]), reducing the fixed costs of employment, and freeing up time for alternative uses. Lower commute times may result in greater work hours for some workers while affecting the labor force participation of others. Alternatively, access to a car may afford greater flexibility in searching for employment. Car owners can cast a wider geographic search net and are not restricted to employment opportunities with time schedules that coincide with transit service schedules.

The impact of lower commute times on employment outcomes can be demonstrated with a simple labor supply model. Figures 1A and 1B depict how access to a car may affect an individual's allocation of time between labor market and non-labor market uses. The time endowment is given by A. Workers convert non-market time into income by supplying time to the labor market at a given wage. Doing so, however, requires commuting to and from work, a time-consuming activity that is not compensated. Hours supplied to the labor market equal the total time endowment, A, less time spent in non-market uses and on commuting. A car affects the budget constraint by reducing commute times from AD to AB. Hence, the budget constraint for a car owner is given by ABC while ADE is the budget constraint without a car.[1]

For a worker with the preferences given in Figure 1A, losing access to a car is equivalent to a reduction in income. This will cause a reduction in consumption of all normal goods. Assuming that free time and market goods (disposable income) are normal, the increase in commute time is absorbed by a decrease in free time (from H^0 to H^1) and a reduction in the consumption of market goods (from I^0 to I^1). The latter change implies a reduction in work hours.

For an individual with relatively strong preferences for non-market time, access to a car may be a deciding factor in the labor force participation decision. This case is depicted in Figure 1B. When facing the budget constraint ABC, this person supplies a small amount of time to the market, generating I^0 in disposable income. At the initial disposable income/non-market time pair, an increase in commute time effectively reduces the wage to zero, causing an overwhelming substitution effect towards non-market uses of time. Here, the individual drops out of the labor market.

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[1]Here we ignore the monetary costs of car ownership and alternative commute modes, subsuming all differences into commute time. To be sure, monetary commute costs are not negligible. However, if total commute costs using alternative transportation measured in foregone time (time spent commuting plus time spent earning the income to cover monetary costs) exceeds comparable costs using a car, the analysis above applies.

Perhaps a more realistic depiction of how car ownership affects employment outcomes is offered by search models that incorporate a spatial dimension. Holzer et. al. [7] present a model of spatial search where owning a car is assumed to reduce per-mile travel costs. Increasing the radius of the geographic area searched increases the probability of receiving an acceptable offer, the expected value of wage offers, and the expected value of the commute distance once a match has been found. Searching workers choose an optimal search radius conditional on per-mile search costs by equating the marginal benefit of search to an exogenously given marginal search cost. The authors show that increasing commute costs per mile decreases the search radius which, in turn, adversely impacts employment outcomes.

Alternatively, car access may impact one's ability to maintain an employment relationship once a match has been made. Given that keeping a job requires showing up regularly and on time, owning a car may lower the likelihood of losing a job (and hence the incidence of joblessness) due to attendance problems. This may be particularly important for workers with complicated commute patterns that are not well-serviced by public transit. For example, for inner-city residents that locate suburban jobs, punctuality may be hindered by public transit commutes that require several transfers and entail a high probability of unscheduled delays. A similar story may apply to workers who must make several intermediate trips between home and work -- e.g., having to drop children off at day care or school.

The existing empirical research implicitly assumes that car ownership is exogenous and regresses employment outcomes on a car ownership indicator and other covariates. Using data from the NLSY, Holzer et. al. [7] find that having access to a car negatively affects the unemployment duration and positively effects wages. In addition, the authors find a larger car-wage effect for black workers than for white workers, a finding consistent with spatial mismatch in urban labor markets. Using a sample of California

AFDC recipients, Ong [10] finds that AFDC recipients with cars are considerably more likely to be employed, and conditional on being employed, work more hours and have higher monthly earnings than recipients without cars.

A major shortcoming of the existing research concerns the lack of attention paid to the issue of causality. A positive correlation between car ownership and employment outcomes is consistent with other hypotheses where cars have no causal effect on labor market prospects. For example, car ownership may be positively correlated with unobserved ability. In this instance, the car dummy variable and the error term in the employment outcome equations will be positively correlated, creating biased estimates of car-ownership effects. Alternatively, car-ownership may in itself be determined by a steady employment history that permits saving and increases access to the capital necessary for making a large purchase. Both scenarios would yield a positive empirical correlations between car ownership and employment outcomes even in the absence of a real causal effect. Below, we propose an empirical strategy that directly addresses this criticism.

3. Empirical Strategy and Data Description

We investigate the effect of car ownership on three employment outcomes: the probability of being employed, weekly work hours, and hourly wages conditional on being employed. To identify the causal effects of car-ownership, we employ several state-level measures of auto-ownership costs as instruments for car-ownership. Using inter-state variation in state-levied gas taxes and automobile insurance premiums, we estimate the model

$$
\begin{aligned}
Employment\ Outcome_{ij} &= \hat{a}'X_{ij} + \tilde{a}Car_{ij} + \varepsilon_{ij} \\
Car_{ij} &= \acute{a}'X_{ij} + \grave{e}Insurance\ Premium_{j} + \ddot{e}Gas\ Tax_{j} + \text{ç}_{ij},
\end{aligned}
\qquad \textbf{(1)}
$$

where i indexes individuals and j indexes states, $Employment\ Outcome_{ij}$ is one of the three outcomes analyzed, X_{ij} is a vector of demographic and human capital variables, Car_{ij} is a dummy variable indicating a car owner, $Insurance\ Premium_j$ is the state average insurance premium, $Gas\ Tax_j$ is the average state gas tax per gallon, â, ã, á, è, ë are parameters, and ε_{ij} and ç_{ij} are normally-distributed error terms. In equation (1), car ownership is modeled empirically with a linear probability model.

To be suitable instruments, insurance premiums and gas taxes must not affect employment outcomes other than through the effects of these variables on the probability of owning a car -- i.e., the instruments must be correlated with Car_{ij}, and be uncorrelated with ε_{ij}. To assess if these conditions hold, one needs to investigate the processes that determine inter-state differences in insurance premiums and gasoline taxes and evaluate whether the determinants of these variables have direct effects on the second-stage employment outcomes.

While there is little research on inter-state variation in insurance premiums, there is a considerable body of literature on the economics of private passenger automobile insurance. This literature may be used to infer sources of geographic variation in insurance premiums. Cummins and Tennyson [2] note that the principal component of insurance premiums consists of the expected payout on each policy. This "pure premium" is determined by two factors: the expected frequency of claims, and the expected severity of claims. Claim frequency and severity may vary geographically for several reasons. For example, there may be variation in the frequency and average length of trips, differences in traffic congestion, or differences in mean demographic characteristics such as age, all factors that may affect state accident rates. In addition, inter-state differences in repair costs and wage levels may affect premiums through payouts for property damages and losses resulting from personal injury.

An alternative source of geographic variation may lie in inter-state differences in regulatory regimes. Under the McCarron-Ferguson Act of 1945, regulation of insurance markets was left to individual states, eventually yielding substantial differences in the institutional structure of auto-insurance markets. For example, more than half of states have some form of rate regulation (Suponcic and Tennyson [16]), ranging from "prior approval systems," where an insurer's rates must be approved by an insurance commissioner, to systems where rates are directly set by a commissioner or an industry rating bureau. Suponcic and Tennyson [16] show that rate regulation is associated with pronounced differences in market structure that are likely to affect insurance premiums. For example, in states with rate regulation, there are fewer firms in the market, fewer multi-state firms, and fewer direct writers or "exclusive dealers" that specialize in auto insurance only and tend to provide coverage at lower costs (Cummins and VanDerhei [3]).

In addition, differences in compensation regimes are also likely to affect expected costs to the insurer and the premiums ultimately paid by consumers. Expected payouts for personal injury awards under no-fault compensation systems are likely to be lower than under traditional tort regimes.[2]

For our purposes, the question of interest is whether the underlying determinants of insurance

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[2]No-fault insurance states require first-party coverage for personal injury and impose thresholds for the severity of personal injuries below which one is ineligible to pursue lawsuits against a third party (Cummins and Tennyson [2]). Given that personal injury payouts constitute nearly 1/3 of the pure premium, the reduction in expected payout via reduced administrative costs and fewer suits is likely to yield a difference in insurance premiums between states with no-fault regimes and states with tort regimes.

premiums in any way directly affect either the probability of employment, usual work hours, or wages. It seems reasonable to argue that inter-state variation in regulatory regimes, trip frequencies and length, and traffic congestion are unlikely to be correlated with the unobservable determinants of employment outcomes.[3] To the extent that such factors are the principal determinants of inter-state variation in premiums, this variable should be a suitable instrument. Inter-state demographic differences, however, in say age, education, gender, race and ethnicity are likely to affect both employment outcomes and insurance premiums. Hence, to the extent possible, our model specifications should include detailed controls for demographic and human capital characteristics.

Geographic variation in local price levels and wages are unlikely to be correlated with unobserved factors that determine employment and hours. These factors, however, are likely to pose problems for our analysis of the effect of car-ownership on wages. If insurance premiums depend on the regional cost of living (via repair costs, for example), and if employers in high cost regions must pay compensating wage differentials, insurance premiums may be correlated with the second-stage wage residuals. This would be particularly true if employers in high-premium states pay wage differentials in compensation for high auto insurance costs. This would impart a downward bias to the 2SLS estimate of car-ownership on wages.[4]

[3]Of course, average trip length by state may be determined in part by the average size of urban areas and urban densities. To the extent that these variables are correlated with racial housing segregation (which may influence state employment rates), insurance premiums may be directly related to our employment outcomes. Below, we control directly for state unemployment rates in an attempt to purge the data of this possibility and to empirically isolate the labor supply effect of car access.

[4]If insurance premiums have a negative effect on the probability of owning a car, the predicted value of car-ownership (instrumented on insurance premiums) will be negatively correlated with a factor (the regional cost of living) that positively effects local wages levels. This yields a negative bias to the 2SLS estimate of the effect of car ownership.

Below, we attempt to directly control for inter-state variation in the cost of living by including estimates of housing costs calculated by the Department of Housing and Urban Development (HUD). However, if insurance premiums depend directly on local wage levels (a fairly plausible proposition), insurance costs will not be a suitable instrument for car-ownership in a log-wage regression. In light of these caveats, the results for wages presented below should be interpreted with extreme caution.

Concerning fuel taxes, state gas taxes are set within a complex system designed to plan, finance, and operate the nation's roads. According to Dunn [5], this system is characterized by two organizing principals: (1) inter-governmental grants flowing from the federal to the state governments, and (2) the trust fund principal. The Federal-Aid Highway and Highway Revenue Acts of 1956 established a federal trust fund financed by the federal gas tax and explicitly earmarked for highway projects. The legislation established a cost-sharing plan where the federal government covers 90 percent of interstate highway construction costs and 50 percent of non-interstate road construction costs. States oversee the construction and maintenance of roads and raise money for the state's share of costs through state gas taxes. The majority of states have exclusive dedication provisions that protect gas tax revenues from competing uses. In particular, 28 states have earmarking provisions in either the state constitution or state law, while eleven others commit receipts to special funds largely devoted to highway projects (Dunn [5]).

This set of institutions suggests that state gas taxes are determined by several factors. First, for federally supported projects, states must raise revenues to meet their matching obligations. Second, given that federal funding is limited by federal gas tax receipts and by the appropriations process,[5] states must raise revenues to fully finance state-initiated road projects for which federal funds are not forthcoming. Hence, combined with the explicit earmarking provisions present in most states, it seems reasonable to argue that the principal determinants of state gas taxes are project financing needs and the availability of federal funds.

For our purposes, the fact that state gas tax revenues are generally separated from state general funds suggests that the fiscal position of individual states is unlikely to affect gas taxes. Hence, demands for state-provided services that fluctuate with the regional business cycle (general assistance, for example) are unlikely to impact gas taxes. Combined with the prominent role of the availability of federal funds in determining state needs, these institutional arrangements suggest that the process determining state gas taxes is independent of state-level differences in employment rates and usual hours worked. Nonetheless, to guard against such a possibility we include state unemployment rates in the specifications of all models estimated below. Concerning wages, to the extent that differences in gas taxes contribute to regional differences in the cost of living (both through direct out of pocket expenses and indirect increases in the costs of goods and services), state gas taxes will not be a valid instrument for car ownership in a log wage equation.

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[5]The initial intentions for the federal highway trust fund were that all receipts be used for highway projects and that the fund would only run small surpluses if any at all. With the passage of balanced budget legislation in 1985, however, pressures to reduce the budget deficit led to appropriations levels below trust fund receipts. Consequently, trust fund surpluses increased substantially after 1985 (Fahey [16]). This period witnessed substantial increases in state gas taxes. Between 1980 and 1990, state gas taxes were increased 136 times and all but three states increased taxes at least once, with average state taxes increasing from 8.6 cents per gallon to 16.8 cents (Dunn [4]). These patterns lend support to the contention that the availability of federal funds is a key determinant of state gas tax levels.

One technical aspect of our empirical strategy that must be addressed concerns the fact that our instruments vary across states yet do not vary across individuals within states. Shore-Sheppard [14] shows that in IV models where the instruments vary between but not within groups, applying the standard estimate of the parameter covariance matrix may lead to faulty inferences.[6] To address this problem, we employ the feasible generalized instrumental variables (FGIV) estimator proposed by Shore-Sheppard. We first estimate the employment outcome models using standard 2SLS and then retrieve the second-stage residuals. We then use these residuals to estimate the within- and between-state variance components and construct the parameter, $è_j = ó^2_{within}/(ó^2_{within} + N_j\, ó^2_{between})$, where *j* indexes states, N_j is the within-state sample size, $ó^2_{within}$ is the within-state variance component, and $ó^2_{between}$ is the between-state variance component of the second-stage residuals. This parameter is then used to quasi-difference all of the endogenous and exogenous variables of the model according to the equation, $y^d_{ij} = y_{ij} - (1 - è^{1/2}_j)\, \mathbf{y}_{.\mathbf{j}}$, where y^d_{ij} is the transformed variable, and $\mathbf{y}_{.\mathbf{j}}$ is the state-level average. We then re-estimate the model applying 2SLS to the quasi-differenced data. This is equivalent to the implicit data transformation applied when estimating random-effects models.[7]

The data for this project are drawn from several sources. Micro data on employment outcomes, car ownership, and basic demographic and human capital characteristics come from the fourth waves of

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[6]Specifically, if the sample is drawn from a population with a grouped structure (or, alternatively stated, if second-stage residuals are correlated within group), the 2SLS estimates of the standard errors may be grossly under-estimated.

[7]Incidentally, the parameter estimates using the FGIV estimator do not differ qualitatively from those using standard 2SLS. However, the correct standard errors are approximately double the estimated standard errors using 2SLS.

the 1992 and 1993 Survey of Income and Program Participation (SIPP).[8] These surveys provide large nationally representative samples of individuals. The fourth wave topical modules of the SIPP collect information on up to three cars per household, including the age of the automobile, the financing status, and the person identifier of the car owners within the household. We use this latter variable to explicitly identify individuals that own a car.[9]

Data on state gasoline taxes as of 1993 and 1994 are assembled by the American Petroleum Institute and are measured in cents per gallon paid at the pump. These estimates include both state excise taxes and sales taxes levied on gasoline purchases. Data on state-level auto insurance premiums come from the National Association of Insurance Commissioners and reflect average expenditures per insured automobile on liability, collision, and comprehensive coverage. We use the SIPP state identifiers to append this data to the micro data sample.[10]

For our linear probability of employment models and weekly hours models, we restrict the sample to civilians, 16 to 65 years of age, with no work-preventing disabilities. For the hours dependent variable, individuals who are not employed are assigned a value of 0. In our log-wage models, we further restrict

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[8]The fourth wave of the 1992 SIPP corresponds to the beginning of 1993 while the fourth wave of the 1993 SIPP corresponds to the beginning of 1994.

[9]The survey provides person numbers for up to two owners. Hence, in a household with two adults and one car where both adults self-identify as being the owner, both adults are coded as owning a car. We experimented with alternative measures of car access, including the presence of a household automobile and the number of cars per adult in each household. The results are qualitatively similar to those presented below using the car-owner variable.

[10]Not all states are individually identified in the SIPP. For some small states, identifiers for state group only are provided. These groups include (1) Vermont and Maine, (2) South Dakota, North Dakota, and Iowa, and (3) Montana, Wyoming, Idaho, and Alaska. For observations residing in these state groups, we assigned weighted group averages of the gas tax and insurance premium figures using state populations for the relevant year as weights.

the sample to wage and salary workers with complete information. Given that the survey collects complete information on all household automobiles only for those households with 3 or fewer cars, we restrict the sample throughout to individuals residing in such households. After taking into account the other sample restrictions, this restriction eliminates approximately 6 percent of the observations.

4. Empirical Results

A. Descriptive Statistics

Table 1 presents baseline estimates of the differences in employment outcomes between those who own cars and those who do not. The table provides mean employment outcomes for the entire sample, individuals by car ownership status, and the differences between those with and without cars. We calculate employment rates and weekly hours using the entire sample while mean log wages pertain to employed individuals only. We present separate calculations for the entire sample and the sample stratified by gender.

There are large, statistically significant differences in employment outcomes for both the entire sample and within gender. Concerning employment rates, there is an approximate 27 percentage point overall difference between individuals with cars and those without, with a slightly higher difference for men (30 percentage points) than for women (24 percentage points). Concerning hours, those with cars work 11 to 16 hours more than those without. Again, the hours differential for men exceeds that for women. Finally, there are substantial differences in the log of hourly earnings between car owners and non-car owners. For the overall sample, the difference in log earnings is approximately .4, with a difference of .5 for males and .3 for females.

To be sure, differences in observable demographic and human capital characteristics are likely to explain much of the differences in employment outcomes between car owners and non-car owners. For example, the descriptive statistics in Table A1 show that individuals without cars are slightly more likely to be female, considerably less likely to be married, and considerably more likely to be black or Hispanic than individuals with cars. In addition, car owners have above average educational attainment, are older than average, and are less likely to be in school. Nonetheless, the baseline difference in employment outcomes presented in Table 1 are substantial, suggesting large upper bounds on the premiums associated with owning a car.

As a first pass evaluation of the strength of our instruments, Figures 2A and 2B present scatter plots of state car ownership rates for 1993 and 1994 against average state auto insurance premiums and state per-gallon gas taxes. We construct state car-ownership rates by averaging the car owner dummy by state and year. The figures also include the predicted relationships between car ownership rates and each of the instruments derived from regressions of the aggregate car-ownership rates on each variable.[11] For both instruments, there are clear negative relationships between the instruments and car-ownership rates. Moreover, the negative slope coefficients are significant at the 0.001 level. For insurance premiums in particular, the data points are compactly distributed around the predicted regression line and the negative relationship does not appear to be driven by outlier states.[12] The data points for the gas tax regressions,

[11]These regressions are weighted by the number of observations for the state-year used to compute the proportion that own cars.

[12]The three observations in the lower right hand corner of Figure 2A include the 1993 and 1994 observations for Washington, D.C. and the 1993 observation for New York. Dropping these observations and re-estimating the car ownership-insurance premium regression yields an estimate of the insurance premium coefficient of -0.000268, with a corresponding t-statistics of -7.15. This is quite close to the estimate provided in Figure 2A.

on the other hand, are less compactly distributed around the regression line. Nonetheless, there are no obvious outliers driving the negative relationship.

To be sure, the first stage relationships evident in the figures may be driven by interstate differences in demographics or economic conditions. To explore this possibility, Table 2 presents means for the variables in our sample after stratifying the data into observations in states with above and below average gas taxes and insurance premiums. The first row presents car-ownership rates while the next three rows present means for our three employment outcomes. Differences in these means are indicative of the reduced form relationship between our instruments and the employment outcomes. The following group of variables provide sample averages for a host of demographic and human capital characteristics that we observe in the data. The next set of variables are state-level characteristics appended to the micro-data sample. Finally, the last set of tabulations indicate the regional distribution of observation for the stratified samples.

The differences in car ownership rates evident in Table 2 are consistent with patterns observed in Figures 2A and 2B. Car ownership rates are higher in low-gas tax states and higher in low-premium states. For employment rates and work hours, residing in a high-tax or high premium state is associated with lower employment rates and lower average work hours. This is clearly consistent with a real effect of car-ownership on these employment outcomes. The reduced-form patterns for wages, however, are the opposite. Wages are higher in high-cost states. This pattern is consistent with the argument that employers in high-cost states must pay compensating differentials to workers.

There are also several differences in the other variables. For example, the unemployment rate at the time of the survey is higher in high-tax relative to low-tax states and in high-premium relative to low-premium states. In addition, low-tax states and low-premium states have lower unionization rates than their high-tax and high-premium counterparts. This pattern is probably due to the disproportionate representation of southern states among the group with below average car ownership costs. An additional pattern of interest concerns differences in housing costs. We appended state-level Fair Market Rent data for 1993 and 1994 calculated by the Department of Housing and Urban Development (HUD) to our microdata set to gauge differences across states in the cost of living.[13] Differences in housing costs between low- and high-tax and low- and high-premium states are quite large and suggest that failure to control for differences in the cost of living will impart a large negative bias to the IV estimates of the effects of car ownership on wages. For example, average Fair Market Rents in high-tax states exceed those in low-tax states by more than $100, while Fair Market Rents in high insurance states exceed those in low-insurance states by more than $170.

The reduced form effects observed for two of the three employment outcomes are suggestive of a real causal effect of car ownership on one's employment prospects. However, the inter-state patterns in several of the other control variables indicate that there may be important differences across states in factors that affect the three employment outcomes and that are correlated with our chosen instruments. In the next section, we present OLS and 2SLS estimates of the impact of car access on employment outcomes that account for the patterns observed in Table 2.

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[13]Fair Market Rents are estimates of the cost of a standard two-bedroom apartment at the 45th percentile of the local rent distribution. These are estimates of gross rent that include the shelter rent plus the cost of all utilities, except telephones. These calculations are used to determine the size of Section 8 housing subsidies and units that are eligible for subsidy under the program. This data is calculated at the PMSA and county level by HUD. We use estimates of corresponding population counts to calculate population weighted state averages. These state-level averages are the figures appended to the microdata.

B. OLS and 2SLS Results

Table 3 presents OLS and 2SLS estimates of the effects of car ownership on each of the three employment outcomes for the entire sample.[14] Specifications of the linear employment probability models and hours models include all variables listed in Table 2 except for the region dummies, the union indicator, and the Fair Market Rent variable (our cost of living proxy). The log-wage regressions include all listed variables less the region dummies. We do not include the region dummy variables in our model specification since this eliminates much of the variation in our instruments. For each outcome, the table first presents the OLS results followed by the second-stage parameters from the 2SLS model. The first-stage, car-ownership regression results are presented in appendix Table A2.

Beginning with the OLS results, controlling for the covariates in Table 2 explains a considerable portion of the unadjusted differences in employment, hours, and wages. Nonetheless, substantial effects remain. After adjusting for observable characteristics there is a 16.8 percentage point difference in the employment rates between car owners and non-car owners, compared with an unadjusted difference from Table 1 of 27 percentage points. In addition, the OLS results indicate that owning a car increases work hours by a bit more than 7 hours and wages by nearly 11 percent, compared with unadjusted effects of 14 hours and 40 percent, respectively. All of the car coefficients from the OLS regressions are significant at the one percent level.

Before turning to the 2SLS results, a brief discussion the first-stage regressions reported in Table A2 is necessary. For all models, each instrument exerts a negative and significant effect on the probability

[14]Recall, the 2SLS results are the results from FGIV estimator outlined above.

of owning a car. In the first-stage results for the employment and hours model, the F-statistics for the test of the joint significance of the two instruments are quite large (95 and 107, respectively).[15] The F-statistic for the comparable joint significance test for the wage model is substantially smaller (7.065). Nonetheless, the effects of gas taxes and insurance premiums are still jointly significant at the .001 level. Hence, the instruments are strongly correlated with the probability of owning a car even after conditioning on a large set of variables.[16]

Turning back to the results in Table 3, the 2SLS estimate of the effect of car-ownership on the probability of being employed (.146) is quite close to the estimates from OLS (.168). Moreover, this point estimate is significant at the 5 percent level. Since we use two instruments in the first-stage regression, we are able to perform a test of the over-identifying restriction, the results of which are reported for each outcome in the last row of the table. For the employment outcome, the test fails to reject the over-identification restriction, suggesting that the strong results for employment are not sensitive to the choice of instruments.

The 2SLS results for hours yield a significant positive effect of car ownership on work hours of approximately 11 hours. While this point estimate is somewhat larger than the OLS estimate of 7.4 hours, the large standard error suggests that this deviation may be due in large part to the imprecision of the point

[15]While the first stage regressions for the employment and hours models are estimated on the same sample, the parameters differ slightly due to differences in the factor used to difference away a portion of the inter-state variation in the data. Recall, the differencing factor, $è_j$, employs estimates of the between- and within-state variance components from the second-stage residuals using initial 2SLS estimates.

[16]Concerning the effects of the other variables in the specification, blacks and Hispanics are considerably less likely to own a car, the probability of owning a car increases with educational attainment, and car ownership increases at a decreasing rate with age.

estimate. Nonetheless, we are able to measure a significant positive hours effect suggesting a real role of access to an automobile on employment and hours. Again, the test of the over-identification restriction fails to reject the restriction at the 10 percent level.

Concerning the 2SLS results for log wages, instrumenting on gas taxes and insurance premiums yields a negative, significant effect of car-ownership on wages. These results are subject to several alternative interpretations. To start, variation in the costs of car-ownership across states may make it necessary for employers in high-cost states to pay wage differentials in compensation for the contribution of these factors to the regional cost of living. To the extent that this is the case, insurance costs and gasoline taxes belong in the second stage regression and hence, the 2SLS model estimated in the final column of Table 3 is mis-specified. Alternatively, these costs may be spuriously correlated with other factors not captured by our measure of inter-state variation in housing costs that affect the local cost of living. Finally, it could be that our estimates are correct and that owning a car reduces wages. This, however, seems highly implausible. We suspect that estimating the wage effect of car ownership requires an alternative identification strategy than that employed here.

With respect to the other variables in the model, most of the point estimates are what one would expect. Blacks are less likely to be employed than whites and earn considerably less per hour. We observe a similar pattern for females and, to a lesser extent, Hispanics. The state unemployment rate exerts a negative effect on the probability of being employed, on hours, and on wages, while state Fair Market Rents have a strong positive effect on log-wages.

Table 4 presents comparable results where we estimate each model separately by gender. Here, we only report the estimated coefficients on the car-owner dummy, the F-statistics for the test of the cumulative significance of the instruments in the first-stage regression, and the test-statistics and P-value for the test of the over-identification restriction. The estimation results for the other background variables do not differ qualitatively from those presented in Table 3 and hence are not reported. We reproduce the results from the pooled sample for ease of comparison.

Again, for both men and women the point estimates of the effect of car ownership on the probability of being employed are similar for the OLS and 2SLS models. The employment effects are considerably larger for women than men and the 2SLS results are significant for women only.[17] For both men and women, we fail to reject the over-identification restriction at the 5 percent level. We observe similar patterns for the hours models. The hours effect of owning a car is slightly larger for women than for men and instrumenting increases the point estimates. Here, the 2SLS results are significant for both men and women. Concerning wages, we find comparable positive effects on the log of hourly wages in the OLS regressions that turn negative (and for women, significant) when we instrument on gasoline taxes and insurance premiums.

In summary, the OLS and 2SLS results both indicate a strong positive effect of car-ownership on the probability of being employed and on work hours that does not appear to be sensitive to the choice of instrument. The results for the employment effect of car ownership indicate substantially larger effects on the labor force participation decisions of women than men. The hours effects are comparable across

[17]While it is difficult to explain this pattern without further information, several potential hypotheses come to mind. To start, the extensive literature estimating labor supply elasticities decisively indicates that male labor supply is considerably less sensitive to net wages than is female labor supply (Killingsworth [9]). A potential extension of this result would be that female labor supply exhibits greater sensitivity to commute costs (a pattern consistent with the findings here). Alternatively, women may, on average, have more complicated commute patterns than men, especially if children are present and if women are more likely to serve as the primary care giver.

gender. Finally, our OLS and 2SLS estimates for the effect of car-ownership on wages are at odds.[18]

6. Conclusion

The results of this paper indicate that having access to a car is an important determinant of labor market outcomes. We find quite strong effects of car ownership on the probability of employment and usual hours worked per week. Moreover, estimates of the effect of cars on these outcomes are comparable in both OLS regressions that ignore the potential endogeneity of car ownership and 2SLS models that employ state-level variation in car operation costs as instruments. Despite our negative 2SLS results for wages, the strong positive effect of cars on wages in the OLS models, coupled with our stated reservations concerning the validity of our instruments in a model of wage determination, suggest that further research on the relationship between car ownership and wages is warranted. Moreover, given the finding that correcting for the endogeneity of car ownership does not appreciably alter our point estimates of the effects on employment and hours (outcomes for which our instruments are better suited), perhaps alternative identification strategies better suited for an analysis of wages would yield comparable results.

[18]We also estimated separate models for sub-samples of the data stratified by potential earnings. We imputed earnings potential for each observation in our sample based on the parameters from a flexibly-specified log-wage regression using data from the 12 Current Population Survey (CPS) Outgoing Rotation Group (ORG) files for 1993. We then stratified the sample into thirds based on position within the imputed earnings distribution. Separate models were then estimated for each sub-sample (see Card [1] and Raphael [13] for detailed discussions of similar imputation procedures). OLS results from this exercise indicate that the impact of car ownership on employment and work hours is larger for low-earnings potential workers relative to high-earnings potential workers. The 2SLS results are similar, yet the car effects by skill group are not significantly different from one another. Interestingly, the impact of the instruments on car ownership are larger for low-earnings potential workers relative to high earnings potential workers. These results are available upon request.

A clear direction for future research would be to assess the degree to which differences in car-ownership rates among distinct groups in the U.S. account for observable differences in employment outcomes. For example, the data in our sample indicate that car ownership rates for Latinos and blacks are slightly more than 20 percentage points lower than those for whites. Moreover, the results from our car-ownership regressions indicate that, even after conditioning on observable human capital characteristics, there is a 16 percentage point difference between the car ownership rates of whites and blacks and a 9 percentage point difference between whites and Latinos.

The finding of a strong employment effect of car access coupled with these large inter-racial/ethnic differences in car ownership rates suggests that car ownership may provide a partial explanation of the persistent racial and ethnic difference in employment ratios and unemployment rates observed in the U.S. This pattern is clearly evident in our data. For example, in OLS linear probability models of the employment outcome, the regression omitting the car ownership variable yields black-white and Latino-white employment rate differentials of 6.9 and 3.3 percentage points, respectively. Adding the car ownership dummy reduces the black-white differential to 4.3 percentage points (a 37 percent reduction) and the Latino-white differential to 1.7 percentage points (a 48 percent reduction). These are large effects with potentially important policy implications.

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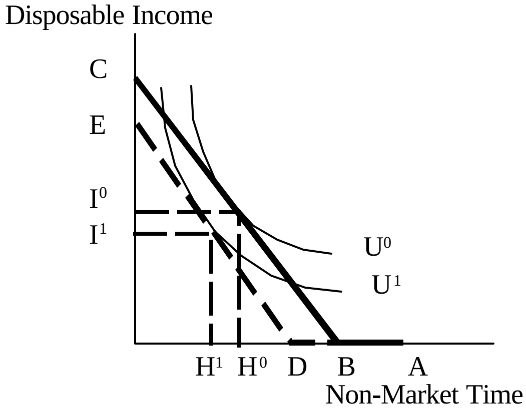
Figure 1

A. Effect on Work Hours



B. Effect on Labor Force Participation

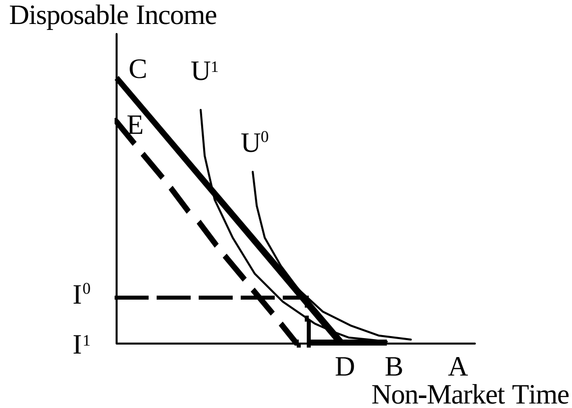


Figure 2A
Scatter Plot of State Car-Ownership Rates Against Average Insurance Premiums

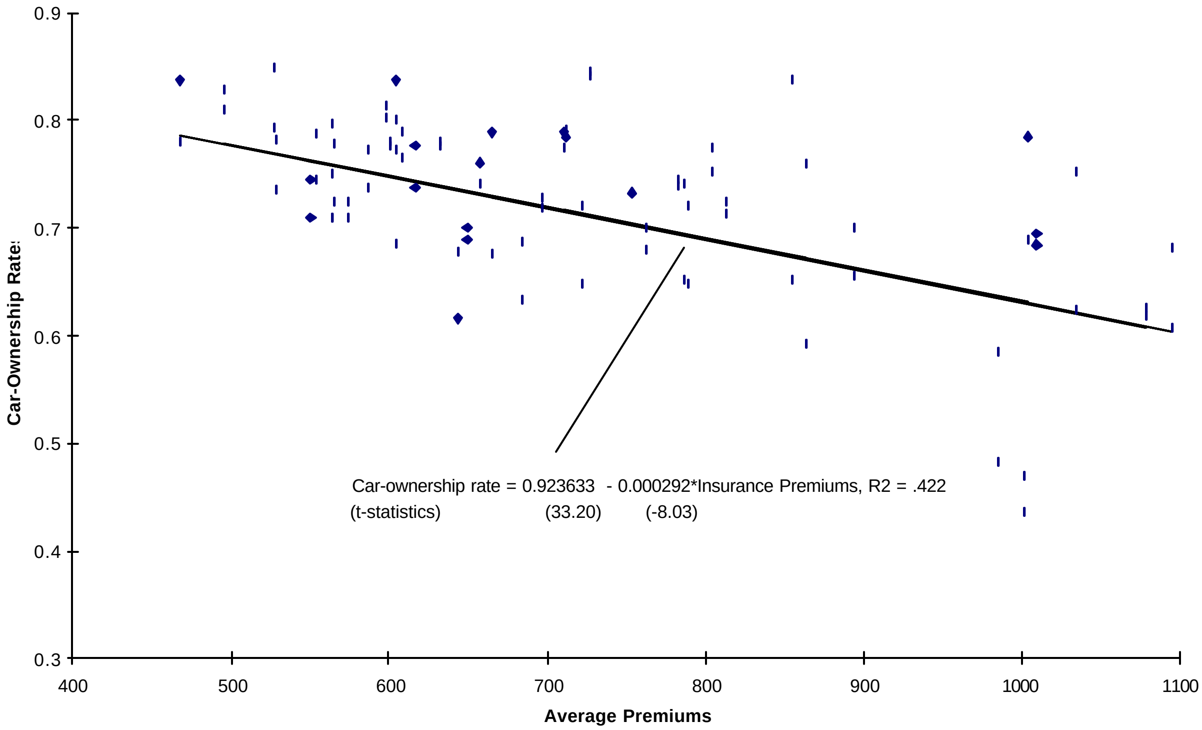


Figure 2B
Scatter Plot of State Car-Ownership Rates Against Per-Gallon Gas Taxes

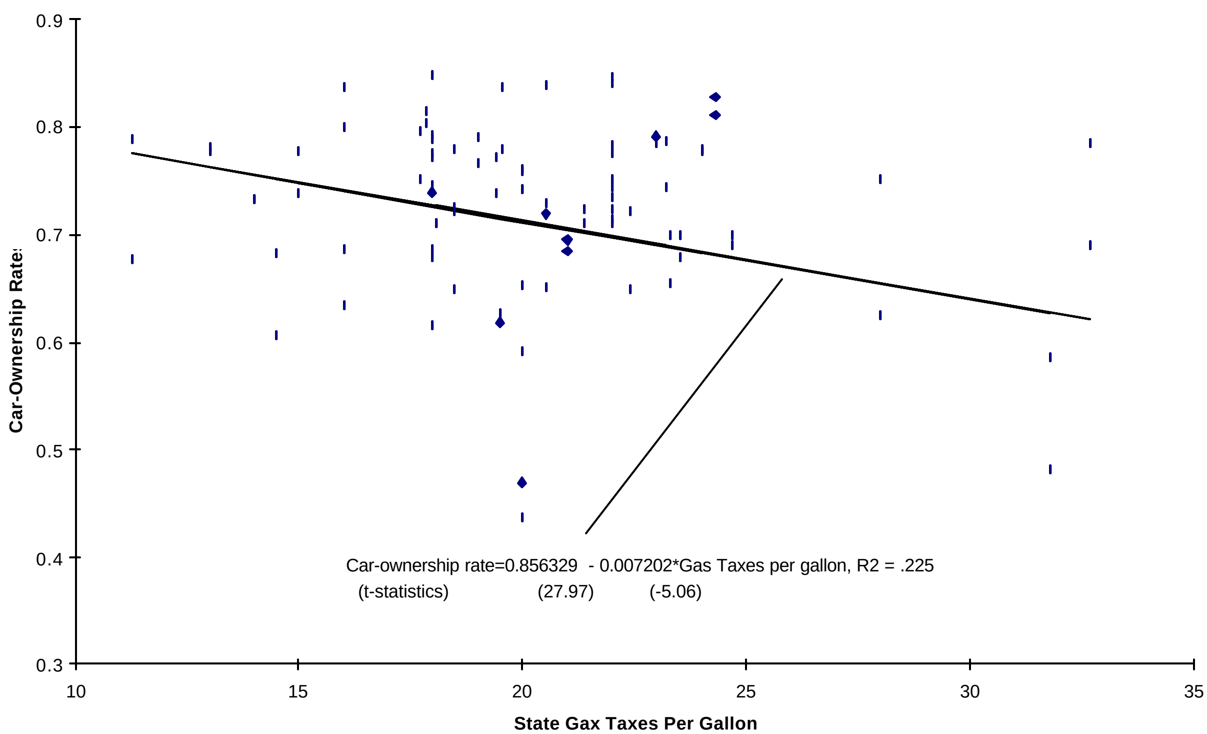


Table 1
Means Employment, Work Hours, and Hourly Wages by Gender and Car Ownership Status

	All	With Cars	Without Cars	Difference
All				
Employed	.715 (.002)	.799 (.002)	.528 (.004)	.271 (.004)
Work Hours	28.687 (.091)	32.887 (.101)	19.265 (.168)	13.622 (.188)
Log Wages[a]	2.278 (.003)	2.368 (.004)	1.981 (.006)	.387 (.007)
N	47,244	33,322	13,922	-
Men				
Employed	.785 (.002)	.875 (.003)	.575 (.006)	.300 (.005)
Work Hours	33.642 (.131)	38.530 (.269)	22.250 (.131)	16.279 (.265)
Log Wages[a]	2.388 (.004)	2.494 (.005)	2.025 (.010)	.469 (.010)
N	21,664	15,697	5,967	-
Women				
Employed	.653 (.003)	.729 (.003)	.488 (.006)	.241 (.006)
Work Hours	24.260 (.120)	27.734 (.140)	16.728 (.209)	11.006 (.251)
Log Wages[a]	2.159 (.004)	2.230 (.005)	1.937 (.009)	.293 (.010)
N	25,580	17,625	7,955	-

Standard errors are in parentheses.
a. Figures are conditional on being employed.

Table 2
Means of Demographic and Background Characteristics for the Total Sample and for the Sample Stratified by the Values of the State Gas Tax and Car Insurance Variables

	Total Sample	Low-Tax States	High-Tax States	Low-Insurance States	High-Insurance States
Car Owners	.691	.710	.674	.729	.654
Employed	.715	.721	.709	.736	.695
Hours	28.687	29.219	28.177	29.391	27.985
Log-Wages[a]	2.278	2.235	2.319	2.234	2.325
Female	.523	.533	.524	.531	.526
Married	.549	.559	.539	.567	.530
Black	.134	.149	.119	.134	.133
Hispanic	.106	.091	.121	.035	.178
Education	13.083	13.043	13.123	13.109	13.058
Age	36.313	36.232	36.390	36.395	36.231
Infant	.100	.101	.099	.095	.104
In School	.157	.159	.157	.150	.163
Union[a]	.149	.120	.176	.136	.162
Unemployment					
Rate	.065	.060	.069	.058	.071
Fair Market	566.066	511.404	618.290	478.560	653.215
Rent[b]	20.983	17.311	24.491	19.903	22.058
Gas Tax[c]	750.183	727.188	772.152	614.964	884.850
Insurance costs					
	.211	.123	.296	.121	.301
North East	.247	.269	.227	.420	.076
North Central	.346	.529	.170	.378	.313
South	.201	.070	.326	.091	.309
West					

a. Conditional on being employed.
b. Rents computed by the Department of Housing and Urban Development for a just standard 2-bedroom apartment at the 45th percentile of the rent distribution.
c. Gas taxes are mean taxes per gallon in cents.

Table 3
OLS and 2SLS Estimates of the Effect of Car Ownership on Employment, Work Hours and Waged

	Employed		Hours		Log-Wages	
	OLS	2SLS	OLS	2SLS	OLS	2SLS
Car-Owner	.168 (.005)	.146 (.076)	7.448 (.205)	10.791 (2.977)	.106 (.006)	-.664 (.363)
Female	-.128 (.004)	-.128 (.004)	-9.443 (.155)	-9.322 (.188)	-.222 (.005)	-.233 (.008)
Married	-.061 (.005)	-.056 (.017)	-2.974 (.187)	-3.716 (.675)	.065 (.005)	.212 (.069)
Black	-.043 (.006)	-.047 (.013)	-1.597 (.250)	-1.134 (.529)	-.088 (.008)	-.191 (.047)
Hispanic	-.022 (.006)	-.023 (.009)	-.369 (.276)	-.160 (.399)	-.102 (.009)	-.165 (.031)
Education	.019 (.001)	.019 (.002)	.966 (.029)	.898 (.067)	.069 (.001)	.080 (.005)
Age	.042 (.001)	.042 (.003)	1.998 (.043)	1.856 (.131)	.056 (.001)	.087 (.014)
Age^2	-.0006 (.00001)	-.0006 (.00003)	-.026 (.001)	-.025 (.001)	-.0006 (.00001)	-.0009 (.0001)
Infant	-.107 (.006)	-.107 (.006)	-4.085 (.268)	-4.118 (.269)	.032 (.009)	.054 (.014)
In School	-.177 (.006)	-.179 (.011)	-12.098 (.263)	-11.711 (.431)	-.115 (.009)	-.194 (.038)
Unemployment	-1.651 (.135)	-1.663 (.237)	-69.292 (5.598)	-65.928 (9.354)	-1.764 (.225)	-2.018 (.498)
Union	-	-	-	-	.178 (.007)	.211 (.018)
Fair Market Rent	-	-	-	-	.0007 (.00002)	.0005 (.0001)
R^2	.198	.193[c]	.295	.317[c]	.369	.284[c]
N	47,244	47,244	47,244	47,244	33,932	33,932
F-Statistic[a] (P-value)	-	94.553 (.0001)	-	107.732 (.0001)	-	7.065 (.0009)
Overid. Test[b] (P-value)	-	1.089 (.310)	-	2.227 (.136)	-	3.232 (.072)

All regressions include a constant. Models are estimated using a grouped error structure.
a. The F-statistics are from tests of the collective significance of the gas-tax and insurance instruments in the first-stage regressions.
b. This is an F-statistic for the over identifying restrictions test of Basmann (1960).
c. Calculated by applying the FGIV parameter estimates to the undifferenced data.

Table 4
OLS and 2SLS Estimates of the Effect of Car Ownership on Employment, Work Hours and Wages, by Gender

	Employed		Hours		Log-Wages	
	OLS	2SLS	OLS	2SLS	OLS	2SLS
A. All						
Car-Owner	.168 (.005)	.146 (.076)	7.448 (.205)	10.791 (2.977)	.106 (.006)	-.664 (.363)
F-Statistic [a] (P-value)	-	94.553 (.0001)	-	107.732 (.0001)	-	7.065 (.0009)
Overid. Test [b] (P-value)	-	1.089 (.310)	-	2.227 (.136)	-	3.232 (.072)
B. Men						
Car-Owner	.142 (.006)	.126 (.095)	6.339 (.298)	9.411 (4.211)	.119 (.010)	-.338 (.377)
F-Statistic [a] (P-value)	-	53.852 (.0001)	-	54.214 (.0001)	-	6.324 (.001)
Overid. Test [b] (P-value)	-	.114 (.735)	-	3.029 (.081)	-	2.523 (.112)
C. Women						
Car-Owner	.174 (.007)	.166 (.086)	7.614 (.273)	11.278 (3.547)	.090 (.009)	-.691 (.360)
F-Statistic [a] (P-value)	-	79.223 (.0001)	-	71.836 (.0001)	-	6.870 (.001)
Overid. Test [b] (P-value)	-	2.852 (.092)	-	.485 (.486)	-	2.271 (.131)

All regressions include a constant and all variables included in the specifications in Table 4. Models are estimated using a grouped error structure.
a. The F-statistics are from tests of the collective significance of the gas-tax and insurance instruments in the first-stage regressions.
b. This is an F-statistic for the over identifying restrictions test of Basmann (1960).

Table A1
Means Demographic Characteristics of Individuals With and Without Cars

	Total Sample	With Cars	Without Cars
Female	.523	.522	.541
Married	.549	.694	.223
Black	.134	.089	.232
Hispanic	.106	.080	.165
Education	13.083	13.461	12.236
Age	36.313	39.497	29.169
Infant	.100	.101	.096
In School	.157	.076	.337

Table A2
First-Stage Results from Regressions of Car Ownership on State-Level Instruments and Demographic Characteristics

	Results from Employment Model	Results from Hours Model	Results from Log-Wage Model
Gas Tax	-.0032 (.0006)	-.0034 (.0005)	-.0023 (.0009)
Insurance Costs	-.0002 (.00002)	-.0002 (.00002)	-.0001 (.00004)
Female	-.0363 (.0034)	-.0363 (.0034)	-.0152 (.0039)
Married	.2169 (.0040)	.2170 (.0040)	.1906 (.0044)
Black	-.1570 (.0056)	-.1575 (.0056)	-.1286 (.0067)
Hispanic	-.0898 (.0062)	-.0894 (.0062)	-.0802 (.0076)
Education	.0208 (.0006)	.0209 (.0006)	.0148 (.0007)
Age	.0421 (.0009)	.0421 (.0009)	.0395 (.0011)
Age^2	-.0004 (.00001)	-.0004 (.00001)	-.0004 (.00001)
Infant	.0090 (.0059)	.0091 (.0059)	.0282 (.0071)
In School	-.1140 (.0058)	-.1139 (.0058)	-.1027 (.0071)
Unemployment	.0573 (.2276)	.0883 (.2172)	-.0252 (.3458)
Union	-	-	.0467 (.0056)
Fair Market Rent	-	-	-.0001 (.00005)
R^2	.332[b]	.333[b]	.120[b]
N	47,244	47,244	33,932
F-Statistic[a] (P-value)	94.553 (.0001)	107.723 (.0001)	7.065 (.0009)

All regression include a constant. All models include random state-year effects.
a. The F-statistics are from tests of the collective significance of the gas-tax and insurance instruments.
b. Calculated by applying the FGIV parameter estimates to the undifferenced data.