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Labor supply in urban China ☆

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Using labor supply responses from 10,560 urban Chinese workers, two-stage least squares estimations identify positive compensated wage effects and negative income effects that are, for the most part, statistically significant. The gross wage effects are mostly positive but they indicate relatively low uncompensated labor supply elasticities. The compensated wage effects are much larger; these may be important in assessing the labor market consequences of reform policies that monetize non-pecuniary benefits. The significance of labor supply responses depends on individual responsibilities within the family; the effects are largest for women and non-household heads. *Journal of Comparative Economics* **31** (4) (2003) 795–817. School of Economics, Georgia Institute of Technology, Atlanta, GA 30332-0615, USA; University of Colorado at Boulder, Campus Box 256, Boulder, CO 80309-0256, USA.

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1. Introduction

Individual labor supply behavior is important for both economic theory and policy evaluation. Verifications of positive compensated wage effects on labor supply confirm an important tenet of microeconomic theory. Demonstrations of negative income effects

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validate the intuition that leisure is a normal good. Furthermore, labor supply responses can be both an objective and a constraint on any policy targeted at individual welfare. Whether labor supply behavior is consistent with standard economic theory in a developing, transitional economy in which the labor market has been partially reformed is particularly important to investigate. The existing literature examines aggregate labor supplies in China; Yang and Zhou (1999) discuss the economic and administrative forces that determine the contemporary allocations of workers to urban and rural sectors. Liang and White (1997), Zhao (1999), and Li and Zahniser (2002) consider the migration of workers from rural to urban areas.

In contrast, the literature on individual labor supply behavior in China is smaller and addresses this issue for rural workers during the era of collective farming. Putterman (1990) examines the relationship between the supply of labor and the proportion of cash in total income. Dong and Dow (1993) estimate the supply of labor to mutual monitoring efforts. Liu (1991) and Burkett and Putterman (1993) analyze the allocation of labor to collective and individual activities. Although these articles are useful to understand the allocation of effort across work opportunities, they do not contain much evidence about the determinants of individual labor provision. Putterman (1990) maintains the hypothesis that uncompensated responses to wage increases are positive. Burkett and Putterman (1993) are unable to discern statistically significant wage effects on annual work days. Moreover, individual labor supply in urban rather than in rural China is analyzed even less.[1]

This paper provides quantitative estimates of individual labor supply behavior using a recent household survey from urban China. Urban labor supply responses to changes in wages and in non-wage income are important because many Chinese reform policies have an impact on wages and non-wage income. For example, the reforms of state-owned enterprises often involve changes in the composition of worker compensation. In many cases, these changes consist of monetizing in-kind compensation by limiting or ending the direct provision of consumption goods and increasing wages so as to allow workers to purchase the same consumption bundle on the open market. Estimate of compensated wage elasticities are necessary to predict the consequences of these reforms on labor supply. In addition, the gross wage elasticity of labor supply has important implications for policy incidence. For the past decade, urban Chinese workers have been subject to mandatory contributions to housing provident funds (Zax, in press). The newly established Chinese pension system requires similar payroll contributions from both employers and workers. The elasticity of labor supply is an important determinant of the incidence of such contributions and the resulting possible losses in employment and efficiency (Li, 2000).

This paper is organized as follows. Section 2 discusses the theory and some empirical models of labor supply. Section 3 describes the data and Section 4 discusses the estimation methods and tests. Section 5 presents the empirical results; Section 6 concludes and discusses policy implications.

[1] Most individual-level research regarding urban Chinese labor markets focuses on the estimation of income functions (Byron and Manaloto, 1990; Knight and Song, 1991; Li, 2003; Liu, 1998; Zax, 1995).

2. Labor supply theory and empirical models

The labor supply behavior of individuals is examined in the context of their families. The collective character of family consumption and investment choices implies that individual labor supply decisions also have a collective component. This is especially important in China because families exert substantial influence over individual choices. The standard unitary model of family labor supply underlies our empirical investigation.[2] In this model, total family utility is determined by total family consumption and by the leisure consumption of each family member. Total family utility is maximized under the family budget constraint to determine the optimal labor supply for each family member.

Assume that the general family utility function is given by $U(C, L_1, L_2, \ldots, L_m, e)$, where C is the composite good of family consumption, L_j is the leisure time of member j, and e represents unobservable family-specific factors that affect the family's utility level. The family budget constraint is given by $C = \sum_j W_j \cdot H_j + V$, where the price of the composite good is normalized to be 1, W_j and H_j are member j's wage rate and working hours, respectively, and V is family property income. In general, the marginal rate of substitution between consumption and leisure for individual j is given as $M_j(\sum W_j \cdot H_j + V, 1 - H_1, 1 - H_2, \ldots, 1 - H_m, e)$. For working member j, working hours are determined by the following equilibrium condition:

$$W_j = M_j\Big(\sum W_j \cdot H_j + V, 1 - H_1, 1 - H_2, \ldots, 1 - H_m, e\Big). \tag{1}$$

This equilibrium condition yields an empirical specification for family member i as follows:

$$H_i = \alpha + \beta \log(W_i) + \gamma\left(\sum_{j\neq i} W_j \cdot H_j + V\right) + \delta' Z_i + e_i, \tag{2}$$

where Z_i is a vector of variables to control for tastes for work, e.g., age and education.

Under this specification, the income elasticity of labor supply is given by

$$\eta_i = \gamma \cdot \left(\sum_{j\neq i} W_j \cdot H_j + V\right)\Big/ H_i. \tag{3}$$

The uncompensated, or gross wage elasticity, is $\zeta_{ui} = \beta/H_i$. Hence, the Slutsky equation gives the pure substitution effect, or compensated wage elasticity, as $\zeta_i = \zeta_{ui} - \gamma \cdot W_i$. Economic theory asserts that this last elasticity should be positive. Wage increases, compensated by reductions in family income to hold utility constant, increase the implicit price of leisure. Therefore, leisure consumption declines and labor supply increases.

If leisure is a normal good, the income effect on labor supply is negative. Uncompensated increases in income increase leisure consumption and correspondingly reduce labor supply. Therefore, the gross wage elasticity can be either negative or positive, depending on the relative magnitudes of the income and substitution effects. At the low wage levels prevailing in urban China, uncompensated wage increases are likely to represent large

[2] Blundell and MaCurdy (1999) provide a convenient summary and, along with Killingsworth (1983), review empirical studies based on this model.

changes in the relative price of leisure. At similarly low levels of income, leisure demand is likely to be low and relatively unresponsive to income. Therefore, the gross wage elasticity is likely to be positive for urban workers in China.

The unitary model leads to a relatively simple empirical specification because it predicts that a family member's labor supply responds only to the aggregate of property income and the labor income of other family members, rather than independently to the individual components. However, it also allows for a rich set of responses because this specification assigns different values for family income to different family members. This model also implies that the correct specification for the labor supply equation contains only aggregate family income, which allows the estimation of a unique income effect. In addition, the hypothesis of equality among the effects of property income and the effects of the labor income of other family members on an individual's labor supply provides an opportunity to test the unitary model. However, these latter implications also have disadvantages. Changes in member j's wage, accompanied by compensating changes in property income, should have no effect on the value of leisure to member i. This implies that wage changes of other family members cannot affect member i's labor supply. In other words, compensated changes in the wages of family member j cannot induce cross-substitution effects for member i.

There are at least two alternative empirical strategies for modeling individual labor supplies in the family context. The male chauvinist model maintains the same assumptions regarding the cross-substitution effects of wage changes. Hence, husbands decide on their labor supplies based solely on own wages and family property income without reference to their wives' labor supply decisions (Killingsworth, 1983). Wives treat husbands' earnings as equivalent to property income (Mroz, 1987). This model is inappropriate for urban China. Most families rely almost equally on the incomes of husbands and wives. The urban female labor force participation rate in 1995 was 79.8%, only slightly lower than the male rate of 83.4%. Moreover, 26.3% of urban Chinese households identified a female as household head (Li and Zax, 2002). Nevertheless, the empirical analysis of Section 4 provides an informal test of this model by including all household income among the determinants of male labor supplies.

The second alternative is a collective model of household labor supply, which is a generalization of both the unitary and male chauvinist models. In this model, individual family members negotiate to a Pareto-optimal allocation of labor supplies and consumption, conditional on the relative bargaining power of individual family members (Chiappori, 1988, 1992). This model allows for asymmetrical treatments of husbands and wives, as in the male chauvinist model, but does not impose such a condition. In the collective model, labor supply responds differently to changes in property income than to changes in wages for other household members because the former are pure income effects but the latter embody two additional effects. First, wage changes may alter the intra-household distribution of bargaining power (Chiappori, 1997). Second, they may change the optimal allocation of family labor to household production activities (Apps and Rees, 1997). Either effect might alter the optimal supplies of individual labor to market activities.

The analysis of Section 4, augmented to distinguish between nonlabor income and the wage income of other household members, provides an informal test of this model.[3]

3. Data and descriptive statistics

The data are from the urban wave of the 1995 China Household Income Project (CHIP-95) survey. This survey covered 6928 urban households and 21,688 individuals located in Anhui, Beijing, Gansu, Guangdong, Henan, Hubei, Jiangsu, Liaoning, Shanxi, Sichuan and Yunnan provinces.[4] Our analysis focuses on the labor supply behavior of working-age individuals and disregards people with ages less than 19 or more than 61 years. In addition, we exclude full-time students, retirees, full-time homemakers, and people with disabilities, injuries or chronic diseases. For our purpose, members of these groups are not considered in the labor force.[5]

Among those that remain in the sample, 97.15% are employed and 2.85% report that they are unemployed or waiting for work. The latter group may include people who choose voluntarily to offer no hours of work, although this sub-group must be quite small because all able-bodied adults intend to work.[6] Therefore, those who are not employed are omitted from the sample. In other words, the question of labor force participation was not important behaviorally for working-age urban Chinese adults in 1995; for these individuals, the amount of labor to supply was the only issue.

The theory of individual labor supply assumes that individuals can vary their work hours continuously. However, many employers have relatively rigid expectations regarding work hours.[7] In conventional analyses of labor supply, this apparent contradiction is reconciled by recognizing that workers choose their work hours indirectly, by choosing their occupations and employers (Killingsworth, 1983, and Blundell and MaCurdy, 1999). For example, work hours will differ for workers who choose to teach in elementary school and for those who choose to work on assembly lines. Therefore, the work hours actually

[3] Formal tests of this model have been restricted to households having both head and spouse present and employed and with no additional workers (Browning et al., 1994 and Fortin and LaCroix, 1997).

[4] This survey was funded by the Ford Foundation, the Asia Development Bank and other institutes, and was conducted under the primary supervision of Zhao Renwei, Li Shi and Carl Riskin. The data have been released to the public at the Inter-University Consortium for Political and Social Research (ICPSR).

[5] Non-participation may not be permanent for this group. As urban Chinese income levels increase, increased levels of support from other family members will continue to discourage participation. However, increasing wages will increase the opportunity cost of non-participation. Hence, this group may merit greater attention in the future.

[6] In more developed countries, the able-bodied working-age population usually contains many more labor market non-participants. As the participation decision also depend on unobservable characteristics, the distribution of these characteristics among those who choose to participate may differ from that among the population as a whole. In this case, ordinary least squares (OLS) estimates would exhibit sample-selectivity bias. As Pencavel (1986) discusses, this bias is less problematic in samples from populations in which larger proportions of individuals work. Urban Chinese labor force participation rates among the able-bodied of working age are virtually 100% so that these biases should be negligible. They may emerge subsequently in urban China if increasing prosperity allows non-participation to become a more common choice.

[7] Lazear (1981) and Kahn and Lang (1992) suggest that restrictions on work hours may arise from differences between instantaneous compensation and marginal products induced by lifetime contracting.

Table 1
Weekly hours by occupational and ownership category

Occupational category	Male household heads	Female household heads	Other males	Other females	Total
Manager, professional or owner					
Average hours	41.62	40.92	41.98	40.95	41.43
Standard deviation	6.72	6.18	6.45	6.40	6.51
Office worker					
Average hours	41.50	40.63	41.19	41.38	41.24
Standard deviation	6.22	5.94	6.37	6.24	6.21
Skilled worker					
Average hours	42.23	41.92	42.67	42.94	42.52
Standard deviation	7.23	6.57	7.58	6.77	7.15
Unskilled worker or other					
Average hours	43.22	42.57	42.97	42.57	42.76
Standard deviation	8.04	7.08	6.78	8.18	7.74
State-owned					
Average hours	41.23	40.85	41.20	41.34	41.20
Standard deviation	6.34	7.00	6.92	5.69	6.41
Local public ownership					
Average hours	42.20	41.23	42.23	41.94	41.97
Standard deviation	6.91	5.94	6.41	7.02	6.68
Urban collective					
Average hours	42.72	42.53	43.25	42.13	42.52
Standard deviation	8.61	6.89	7.87	8.43	8.13
Other					
Average hours	42.97	40.75	46.94	45.52	45.17
Standard deviation	6.89	9.38	7.67	6.78	7.46

Notes. 1. The manager, professional or owner category consists of individuals identifying themselves as division head in institution, head of institution, professional or technical worker, owner of private or individual enterprise or owner and manager of private enterprise. 2. The other category for ownership consists of private enterprise, including partnership, self-employed business/individual enterprise, Sino–foreign joint venture, foreign owned, township and village enterprises and a residual category.

observed for any worker can be construed as reflecting labor supply choices only if workers have some freedom to choose their jobs. In pre-reform urban China, this opportunity was almost entirely absent, although it has become increasingly widespread as the reforms have progressed. By 1995, individuals were not only able to choose their occupations but also to move across employers, both within and across ownership sectors. For example, among those in the CHIP-95 data with six to ten years of work experience, approximately 22% had changed jobs at least once (Li and Zax, 2002).

The CHIP-95 data demonstrate that variation in work hours exists across occupations and ownership types. As shown in Table 1, typical work weeks are longest for workers of lower skill and occupational status. Average weekly work hours for unskilled or other workers exceed those for all other occupations. On average, unskilled workers work about one and one-quarter more hours than do managers and professional workers per week. Variations in work hours across ownership types are observed as average weekly work hours are smallest for workers in state-owned enterprises. Among enterprises owned by local government, the average weekly work hours are approximately three-quarters of an

Table 2
Work hours and wages

	Male household heads	Female household heads	Other males	Other females	Total
Average hours per week	41.93	41.35	42.20	41.94	41.91
Standard deviation	6.942	6.426	6.850	7.103	6.90
Average hourly wage	3.488	3.040	2.835	2.548	2.971
Standard deviation	2.385	1.781	2.005	1.868	2.088
Correlation between wage and weekly hours	−0.3951	−0.4084	−0.3668	−0.3360	−0.3652
p-value	0.0001	0.0001	0.0001	0.0001	0.0001
Observations	3167	1627	2407	3359	10,560

hour more than among those owned by higher levels of government. In the non-public sector, including all forms of joint ventures and private ownership, the average work week is generally the longest. Hence, the urban Chinese labor market presents workers with a fairly wide choice of work hours. The evidence that work hours vary across employers and that workers are mobile across employers suggests that preferences regarding labor supply in urban China may be expressed through the choice of occupation and employer.[8] Therefore, as in conventional labor supply studies, the analysis of Section 5 treats work hours as a continuous choice variable.[9]

Our sample consists of 10,560 working individuals in urban China who are between 19 and 61 years old.[10] Table 2 presents averages for weekly hours worked. The table indicates that, first, the typical work week is longer than in most developed countries, at approximately 42 hours. Additionally, typical work weeks are surprisingly uniform across individuals having different household responsibilities. Two further observations emerge from the standard deviations of weekly work hours. First, the range of variation in weekly work hours is similar across individuals having different household responsibilities. Second, this variation is relatively uniform across all categories and modest in magnitude. The coefficient of variation for weekly work hours varies from 0.155 to 0.169 across the four categories.

According to labor supply theory, the primary determinant of work hours is the price of leisure, which is the wage. From Table 2, this price is highest for male household heads. On average, their hourly wage exceeds that of workers in all other categories of household

8 In addition, workers can vary total labor supply by absences from work, overtime work or second jobs (Conway and Kimmel, 1998). The CHIPS-95 data do not report overtime work, and only fragmentary information is available regarding sick leaves and moonlighting.

9 Dickens and Lundberg (1993), Kahn and Lang (1992), Makunnas and Pudney (1990) and Tummers and Woittiez (1991) offer labor supply estimation strategies that incorporate demand-side restrictions on hours worked explicitly. We adopt the simpler strategy of treating work hours as a continuous variable to focus on the novelties of Chinese data and institutions.

10 We exclude individuals with incomplete data, those working an average of more than ten hours a day and seven days a week, and those reporting imputed hourly wages of less than 0.15 yuan. This last exclusion is based on comparisons with the average national wage in 1995 of 2.20 yuan/hour (China Statistical Publishing House, 1997, p. 40).

responsibility by at least approximately one-half of a yuan.[11] Female household heads earn, on average, approximately one-half yuan more than do female non-heads. In other words, regardless of sex, headship appears to be associated with a wage premium of approximately 20%.[12] The empirical correlations between weekly hours and hourly wages are uniformly negative, sizeable, and very significant, regardless of household responsibility. Although these correlations incorporate a wide array of direct and indirect structural relationships, compensated labor supply responses to wage changes are clearly not dominant. The pure effect of an increase in the price of leisure on leisure consumption and labor supply should be negative and positive, respectively. If it is present, this effect is dominated by income or other effects that reduce weekly work hours when wages increase.

Of course, Table 2 is only descriptive. Therefore, these implications are no more than suggestive. Models that account simultaneously for all important determinants of work hours would be necessary to provide more compelling evidence. The next section discusses their specification.

4. Specifications for regressions

In estimating the empirical model, the principal challenge is the possibility that the two explanatory variables of interest, own wages and family income, are endogenous to the determination of hours of work. Wages and hours worked are likely to be mutually dependent, for reasons of both theory and measurement (Fortin and LaCroix, 1997). Theoretically, hours and wages are chosen jointly so that they may both be consequences of unmeasured individual characteristics, such as motivation and ambition. These characteristics would be embodied in both the wage and the disturbance term of a regression equation predicting work hours. Hence, the wage and disturbance terms would be correlated. Empirically, hourly wages are calculated as the ratio of annual income to estimated annual work hours. Therefore, any measurement error in work hours would appear in the wage measure, which would also lead to correlations between wages and the disturbance term in the regression equation.

Family income for each individual includes labor income received by other family members and all non-labor income received by the family. Labor income received by other family members may be endogenous because member labor supplies are determined jointly

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[11] This wage measure is the ratio of total annual labor income to annual work hours, estimated as the product of average daily work hours and average work days per week times 52 weeks. On average, estimates of hourly wages derived from annual earnings and hours appear to be substantially more accurate than estimates derived from reports of usual earnings and hours or earnings and hours during the most recent pay period in the United States. Nevertheless, individual errors can be substantial (Bound et al., 1990; Bound and Krueger, 1991; Rodgers et al., 1993). There appear to be no analogous validation studies for other countries and certainly not for China.

[12] A substantial literature documents marital wage premiums for American men (Korenman and Neumark, 1991). Most explanations for this premium imply that it is the cause, rather than the consequence, of headship. At the same time, marriage is associated with reduced wages for American women. The comparisons in Table 2 suggest that it may be worthwhile to explore the effects of headship, as opposed to marriage, on American female wages.

in either the unitary or collective models of household decision-making. Components of family income may also be endogenous. For example, transfer payments from either the government or from individuals outside the family could depend directly on labor supply.[13]

Instrumental variables estimation is a potential solution for such simultaneity bias. Regarding the own-wage, the existing literature relies mainly on worker characteristics that are ordinarily excluded from the hours equation, such as higher order terms in age, experience or education, for instruments (Mroz, 1987; Sahn and Alderman, 1996; Fortin and LaCroix, 1997). However, additional instruments may be available from the demand side of the labor market. Variables describing the general structure of the local labor market may be related to the wage received by an individual worker but unrelated to the idiosyncratic component of that worker's individual labor supply choice. We use a demand-based instrument that represents a novel identification strategy in labor supply estimation.

The six instruments that we use for the own-wage include both individual and market level variables. Two of these, namely, the square of work experience and the interaction of work experience and age, are conventional instruments in labor supply estimations and measure additional human capital characteristics. Two variables, namely, total prior months of full-time and part-time on-the-job training, should be related to the wage because of human capital accumulation but may be exogenous to current labor supply choices. One measure of workplace conditions, namely, an index of extreme temperatures, may be related to wages due to compensating differentials.[14] The last variable, namely, the proportion of county employment in the non-public ownership sector, should be related to individual wages because of its influence on the wage structure in the local labor market.

Family income may be heterogeneous within a household and its composition may vary substantially across households. Therefore, theories that predict revenue amounts from any individual source are unlikely to yield instruments that are effective for all households. However, time series data with lagged values may provide instruments that are statistically appropriate. The CHIP-95 data include measures of individual incomes in previous years. In most households, the incomes that accrue to other household members form the bulk of family income for each household member in each year. The sum of these incomes may be correlated with family income in the current year but not with the disturbance term in the equation for current work hours, if the idiosyncratic components of individual labor supply are uncorrelated across years. Thus, we use family income for 1993 and 1994 as the instruments for current family income in our analysis.[15]

In instrumental variables estimation, a priori arguments cannot eliminate entirely suspicions that some instruments might be correlated with elements of the disturbance

[13] Sahn and Alderman (1996) treat own wages as endogenous but disregard these issues and treat the wages of other family members as exogenous.

[14] Wage differentials that compensate for workplace disamenities should, by definition, have effects on labor supply only if compensation is not exact.

[15] The survey requested incomes for all years beginning in 1990 but information prior to 1993 was lacking. Basmann tests reject the hypothesis that the sum of incomes of other household members was orthogonal to current work hours when family income in 1992 is included. This may be caused by the accumulation of measurement errors through changes in family composition or in labor force participation.

term. Empirical tests on overidentifying restrictions address these suspicions. The Basmann test (Basmann, 1960) examines the validity of the orthogonality condition between instruments and the errors when the number of instruments exceeds the number of endogenous regressors, given that a subset of the instruments are valid and identify the model.[16] In our models, if any two instruments are valid, the overidentification test assesses the validity of the excess instruments and the assumption that all other regressors are exogenous. In our analysis, the test results fail to reject the overidentifying restrictions for all but one of the models. Given the fairly large number of additional instruments, such a result offers strong support for the model specifications and for the instruments.

In addition to potential endogeneity, measures of household income may be distorted in two other dimensions. First, variations in household size alter the available income per capita. The regressions in Section 5 control for these variations by using the number of household members.[17] Second, current income may contain substantial transitory components. For example, in a transition economy like China, flows of asset incomes are likely to change over time and be reflected in asset valuations. Therefore, asset valuations may contain additional information regarding the household's intertemporal budget constraint that would affect individual labor supplies. For this reason, the regressions in Section 5 contain measures of household assets. For many workers in China, housing constitutes the single largest asset. The implicit value of housing appears in three variables; namely, dummy variables for owned private housing and publicly-owned housing and non-labor income, which includes imputed rents as the difference between estimated market rents and actual rents.[18]

According to labor supply theory, hours worked are determined by individual tastes for work in addition to wages and non-labor income. Therefore, the regressions control for several exogenous proxies related to these tastes, including sex and household headship. Successive estimations present increasingly detailed stratifications by these variables. All regressions contain variables measuring completed education, work experience and age. Levels of human capital, proxied by the former, may affect labor supply through their effects on home productivity (DaVanzo et al., 1976). In addition, among workers receiving

[16] Hansen (1982) provides a more general version of this test using the generalized method of moment (GMM) framework. Our estimations employ eight instrumental variables for two potentially endogenous explanatory variables. Generally, instrumental variables (IV) estimation becomes asymptotically more efficient as the number of instruments increases. In order for the IV estimator to have a mean and variance, the number of instruments should be equal to or larger than the number of endogenous variables plus two (Kinal, 1980). Increases in the number of instruments also have the disadvantage that they increase the finite sample biases of IV estimators (Davidson and MacKinnon, 1993). However, this is a minor concern given our degrees of freedom available.

[17] Mroz (1987) and Xie (1997) conclude that the number of children in a household is not endogenous with respect to either male or female labor supply in US data.

[18] Accounting for this asset is complicated because, in 1995, virtually no urban housing was exchanged on a market basis (Zax, in press). Many families in owner-occupied housing provide an estimated market value for their dwelling unit. However, families in government-owned housing may have an implicit property right to their housing, the value of which is unknown. Our specification attempts to differentiate between different ownership claims without relying on a uniform valuation. Virtually all families reported an estimate of the market rent for their dwelling. Imputed rents are equal to this estimate minus actual rents; this calculation assumes that the latter is zero for owner-occupants. We disregard possible interest costs because these are not reported and few owner-occupant families report the indebtedness that would be associated with house purchases.

the same wage, differences in levels of human capital may indicate differences in tastes for work. Experience may capture the evolution of these tastes over the course of a career (Nakamura and Nakamura, 1981). Age may capture life-cycle effects and variations in household responsibilities within headship categories.

Health affects an individual's capacity for work and therefore, it has an impact on labor supply. As proxies for this variable, the regressions include continuous variables measuring health care expenditures by each individual and by the state on the individual's behalf in addition to the average number of cigarettes smoked per day. Finally, the exogenous variables include two individual characteristics that may be institutionally relevant. Communist party membership may have either a positive or negative effect on labor supply. Ideology may exhort individuals to exceptionally high work effort; alternatively, privilege may protect individuals from sanctions for minimal efforts. Workers who were rusticated in their youth, that is, sent to rural areas during the Cultural Revolution, usually suffered in both the quality and the quantity of their education.[19] Moreover, this experience may have changed their tastes regarding work and leisure so that these individuals may have distinctive labor supply patterns. The inclusion of this variable offers an opportunity to examine the nature of the reintegration of these individuals into urban Chinese society.[20]

5. Estimation results

Table 3 presents definitions and descriptive statistics for all variables, including the instruments. The dependent variable is annual hours of work, estimated as the product of average daily work hours and average work days per week multiplied by 52. This rescaling matches the income variables, which are in annual terms. This estimate omits vacation time, other absences from work and overtime. Although unfortunate, these omissions are common, especially for the latter two (Killingsworth, 1983).[21] Consequently, measured work hours are usually error-ridden approximations of actual work hours. From an econometric perspective, these errors are not problematic because they arise in the dependent variable. According to conventional analysis of errors in variables, random measurement error in the dependent variable does not bias estimations. The assumption

[19] Those young people were required to move to rural areas during the "Up to the mountains and down to the countryside" campaign of the Cultural Revolution. Rustication was based on political judgments; therefore, it is unlikely to be related to individual preferences regarding work. Thus, this variable should not be correlated with unobserved factors that affect the individual's labor supply.

[20] Demand side variables such as occupation, industry and ownership type are not included in conventional labor supply models. As discussed in Section 3, workers choose these attributes simultaneously with their choice of labor supply. Therefore, they are not exogenous determinants of labor supply.

[21] The CHIP-95 data do not record paid vacations. As of 1995, Chinese labor law stipulates that all workers with at least one year of seniority are entitled to paid vacation. However, the law does not require any specific amount of vacation, nor does it distinguish between industries or ownership type. No known source, either in English or Chinese, studies variations in actual paid vacations among urban workers. Some labor experts in China believe that paid vacation time varies by employer rather than by industry or ownership type. In this case, this variation would appear as a random measurement error of the dependent variable in the error terms of the specifications.

Table 3
Descriptive statistics

Variable	Definition	Mean	Std. dev.	Min	Max
Hours worked	Annual working hours, hours/year	2179.30	358.81	208	3432
Wage	Wage rate, measured by yuan/hour	2.97	2.09	0.15	39.32
Non-wage income	Non-wage family income, in thousand yuan	12.33	25.77	0	974.50
Home assets	Household assets, in thousand yuan	23.24	35.52	−379	820
Age	Age	38.72	9.50	19	61
Years worked	Actual years of work experience	19.58	9.58	1	47
Party affiliation	1 if Communist party member, 0 otherwise	0.26	0.44	0	1
Family size	Number of family members	3.32	0.79	1	8
Housing ownership, public	1 if housing is publicly owned, 0 otherwise	0.55	0.50	0	1
Housing ownership, private	1 if housing is privately owned, 0 otherwise	0.44	0.50	0	1
Healthcare cost, individual	Individual expenditure on medical care (yuan)	111.91	314.54	0	7644
Healthcare cost, state	State paid medical expenses for a family member (yuan)	306.09	1153.37	0	40,000
Rusticated youth	1 if the member was sent as youth to the countryside, 0 otherwise	0.22	0.41	0	1
Cigarettes per day	Number of cigarettes smoked per day	4.32	7.76	0	60
Education level, university	1 if member attended college or above,	0.08	0.27	0	1
	0 otherwise	0.08	0.27	0	1
Education level, professional	1 if member's highest level of education was attendance at professional school, 0 otherwise	0.16	0.36	0	1
Education level, technical	1 if member's highest level of education was attendance at middle level professional or technical school, 0 otherwise	0.17	0.38	0	1
Education level, upper middle	1 if member's highest level of education was attendance at upper middle school, 0 otherwise	0.24	0.43	0	1
Education level, lower middle	1 if member's highest level of education was attendance at lower middle school, 0 otherwise	0.30	0.46	0	1
Gender of member	1 male, 0 female	0.53	0.50	0	1
Status in family, head	1 household head, 0 otherwise	0.45	0.50	0	1
Status in family, spouse	1 spouse of household head, 0 otherwise	0.39	0.49	0	1
Status in family, child	1 child of household head, 0 otherwise	0.15	0.36	0	1
Status in family, parent	1 parents of household head, 0 otherwise	0.0019	0.043	0	1
Full time training	Months of full-time on-the-job training	1.32	4.65	0	70
Part time training	Months of part-time job training	0.90	4.05	0	65
Work temperature	1 if work under high temperature or low temperature, 0 otherwise	0.05	0.22	0	1
Non-public employment	Percentage of employment in non-public sectors in the county	0.020	0.029	0	0.23
1993 Non-wage family income	Non-wage family income in 1993, in thousand yuan	5.22	4.04	0	77.1
1994 Non-wage family income	Non-wage family income in 1994, in thousand yuan	6.29	4.81	0	95

Notes. 1. The total number of observations is 10,560. 2. The omitted category of housing ownership is rented from private owner and other. 3. The omitted category of education is elementary school and below. 4. Member status-child-includes children and children in law; member status-parent-includes parents and parents in law. 5. The omitted category of family membership includes all other members live in the household.

of random measurement error is maintained in our analysis following most studies of labor supply.

Table 4 presents the two-stage least squares regression specifications described in the previous section. The ordinary least squares (OLS) estimation, which is reported in Table 5, yields coefficients for family income and family assets that are positive and

Table 4
Overall labor supply and labor supply by gender

Variable	2SLS (all) (instrument wage)	2SLS (all) (instrument wage & income)	2SLS (men) (instrument wage & income)	2SLS (women) (instrument wage & income)
Wage (in log term)	36.83*	118.63**	113.91	161.18**
	(1.65)	(2.60)	(1.38)	(2.81)
Non-wage income	−0.33**	−5.25**	−5.37*	−6.52**
	(−3.93)	(−3.02)	(−1.78)	(−2.84)
Home assets	−0.049	0.15	0.21	0.080
	(−0.38)	(1.14)	(1.29)	(0.32)
Age	−1.78	−9.83*	−10.009	−23.76*
	(−0.44)	(−1.76)	(−1.01)	(−2.55)
Age squared	−0.023	0.082	0.055	0.29**
	(−0.49)	(1.23)	(0.49)	(2.46)
Years worked	1.81*	0.60	2.058	−0.17
	(1.64)	(0.47)	(1.14)	(−0.10)
Party affiliation	10.34	4.88	24.081*	−36.52**
	(1.17)	(0.48)	(1.88)	(−2.19)
Family size	26.43**	44.75**	42.66**	52.98**
	(5.15)	(4.96)	(2.96)	(4.19)
Home ownership, public	0.31	7.97	−22.75	46.54
	(0.01)	(0.26)	(−0.55)	(0.97)
Home ownership, private	33.41	54.51*	27.99	90.44*
	(1.13)	(1.71)	(0.63)	(1.84)
Healthcare cost, individual	0.023	0.028*	0.026	0.035*
	(1.63)	(1.81)	(1.11)	(1.77)
Healthcare cost, state	−0.0090**	−0.010**	−0.010**	−0.012**
	(−2.79)	(−2.99)	(−2.12)	(−2.51)
Rusticated youth	−13.93	−13.24	5.73	−31.73**
	(−1.50)	(−1.31)	(0.36)	(−2.10)
Cigarettes per day	0.86*	0.82	0.90	0.38
	(1.59)	(1.36)	(1.40)	(0.19)
Education level, university	−146.86**	−174.09**	−178.80**	−180.52**
	(−5.74)	(−5.72)	(−4.10)	(−3.97)
Education level, professional	−128.69**	−148.75**	−150.59**	−160.70**
	(−5.82)	(−5.77)	(−4.18)	(−4.20)
Education level, technical	−110.51**	−130.39**	−141.16**	−126.96**
	(−5.06)	(−5.12)	(−3.80)	(−3.41)
Education level, upper middle	−67.47**	−82.01**	−110.70**	−50.42
	(−3.25)	(−3.49)	(−3.27)	(−1.50)
Education level, lower middle	−41.36**	−51.81**	−72.43**	−32.59
	(2.06)	(−2.38)	(−2.37)	(−1.02)
Gender of member	16.27*	4.76		
	(1.90)	(0.46)		

(*continued on next page*)

Table 4 (*Continued*)

Variable	2SLS (all) (instrument wage)	2SLS (all) (instrument wage & income)	2SLS (men) (instrument wage & income)	2SLS (women) (instrument wage & income)
Member status in family, head	-161.88^{**}	-223.39^{**}	-183.07^{*}	-263.02^{**}
	(−2.45)	(−3.14)	(−1.56)	(−2.73)
Member status in family, spouse	-151.4^{**}	-198.08^{**}	-172.67^{*}	-223.57^{**}
	(−2.30)	(−2.88)	(−1.54)	(−2.36)
Member status in family, child	-141.70^{**}	-184.40^{**}	−144.78	-223.27^{**}
	(−2.18)	(−2.80)	(−1.38)	(−2.47)
Member status in family, parent	-185.37^{*}	15.64	425.54	-318.63^{*}
	(−1.79)	(0.06)	(0.70)	(−1.92)
Observations	10560	10560	5574	4986
F-statistic	(24, 10535) = 9.46	(24, 10535) = 8.03	(23, 5550) = 4.57	(23, 4962) = 4.46
Prob > F	0.00	0.00	0.00	0.00
Overidentifying restriction test (F-stat)	(5, 10530) = 1.69	(6, 10529) = 1.65	(6, 10529) = 1.66	(6, 10529) = 1.66
Prob > F	0.13	0.13	0.13	0.13
Hausman test (F-stat)	(1, 10534) = 205.66	(2, 10533) = 105.96	(2, 5548) = 55.75	(2, 4960) = 64.49
Prob > F	0.00	0.00	0.00	0.00
Calculated elasticity				
Income elasticity	−0.002	−0.030	−0.030	−0.03
Gross wage elasticity	0.017	0.054	0.052	0.06
Compensated wage elasticity	0.997	15.65	17.29	17.74

Notes. 1 The constant term is not reported. All standard errors are heteroskedasticity robust, and t-statistics are in parentheses. 2. The elasticity is calculated based on the relevant sample averages of working hours, non-wage family income, and wages.

* Significance at the 10% level.

** Significance at the 5% level.

Table 5
Other estimation results

Variable	OLS (all)	2SLS (all head) (instrument wage and income)	2SLS (all non-head) (instrument wage & income)
Wage (in log term)	-231.95^{**}	-11.25^{*}	193.42^{**}
	(−26.69)	(−0.16)	(3.70)
Non-wage income	0.28^{*}	-2.31^{**}	-6.23^{**}
	(1.92)	(−0.91)	(−3.15)
Home assets	0.95^{**}	0.33	−0.034
	(7.63)	(1.80)	(−0.18)
Age	14.29^{**}	−3.51	−18.20
	(3.95)	(−0.42)	(−2.53)
Age squared	-0.25^{**}	−0.0094	0.20
	(−5.81)	(−0.10)	(2.20)
Years worked	6.90^{**}	2.62	−1.073
	(7.01)	(1.54)	(−0.60)
Party affiliation	29.90^{**}	19.21	−6.60
	(3.77)	(1.51)	(−0.40)

(*continued on next page*)

Table 5 (*Continued*)

Variable	OLS (all)	2SLS (all head) (instrument wage and income)	2SLS (all non-head) (instrument wage & income)
Family size	12.87**	26.06**	59.39*
	(2.77)	(2.05)	(5.32)
Home ownership, public	9.97	−20.61	39.11
	(0.36)	(−0.47)	(0.95)
Home ownership, private	33.99	3.57	100.80*
	(1.24)	(0.08)	(2.35)
Healthcare cost, individual	0.022*	0.043	0.013*
	(1.76)	(2.43)	(0.56)
Healthcare cost, state	−0.00178	−0.0068**	−0.014**
	(−0.66)	(−1.77)	(−2.481)
Educated youth	−11.99	−7.94	−26.40
	(−1.45)	(−0.60)	(−1.70)
Cigarettes per day	0.77	0.82*	1.29
	(1.60)	(1.26)	(1.44)
Education level, university	−9.25	−150.98**	−190.27**
	(−0.42)	(−3.71)	(−4.27)
Education level, professional	−25.56	−145.91**	−151.50**
	(−1.30)	(−4.47)	(−3.95)
Education level, technical	−16.79	−129.05**	−126.35**
	(−0.86)	(−3.86)	(−3.29)
Education level, upper middle	−7.65	−118.40**	−44.42**
	(−0.40)	(−3.81)	(−1.28)
Education level, lower middle	−7.80	−89.58**	−12.96**
	(−0.42)	(3.17)	(−0.40)
Sex of member	42.24**		
	(5.47)		
Member status in family, head	−20.02		
	(−0.25)		
Member status in family, spouse	−37.28		
	(−0.460)		
Member status in family, child	−44.90		
	(−0.56)		
Member status in family, parent	−179.98*		
	(−1.71)		
Observations	10560	4794	5766
F-statistic	(24, 10535) = 43.16	(19, 4774) = 4.58	(19, 5746) = 6.60
Prob > F	0.00	0.00	0.00

Notes. The constant term is not reported. All standard errors are heteroskedasticity robust, and t-statistics are in parentheses.

* Significance at the 10% level.

** Significance at the 5% level.

significant. Taken literally, they imply implausibly that leisure is an inferior good. Both of these coefficients become negative when the own wage variable is treated as endogenous. Furthermore, the estimated gross wage effect switches from negative to positive.

In column 2 of Table 4, the treatment of both wage and family income as endogenous yields estimates of the gross wage and family income effects that continue to have the correct signs, are much larger in absolute value, and are both significant at better than the 1% level. The Basmann test does not reject the overidentifying instrumental restrictions and the Hausman test rejects the OLS estimation, which implies that the OLS estimates are biased because of endogeneity. [22] Taken together with the discussion in Section 4, these results justify our reliance on this specification in the remaining estimations.

Although significant, the magnitudes of the wage and family income effects are relatively small. From column 2 of Table 4, at the overall sample average values of approximately 2179 hours of work per year, wages of 2.97 yuan per hour and family income of 12,330 yuan, the uncompensated elasticity of annual work hours with respect to wages is 0.054. The elasticity of annual work hours with respect to family income is -0.03.[23] However, the implied compensated elasticity of annual work hours with respect to the wage is much larger, at 15.65. In other words, wage increases that are compensated by corresponding reductions in other income elicit substantial increases in work hours.[24]

Compensated changes of this sort are not relevant for most policy purposes. However, the reduction of labor entitlements, i.e., monetary and non-pecuniary benefits that are not related to productivity, is a component of urban Chinese labor market reforms. If these entitlements are replaced by higher wages, the net effect would be to create the compensated wage change to which the above elasticity applies. For example, urban housing reforms require work units to raise rents on worker housing units that are compensated for by commensurate wage increases (Zax, in press). The net effect of these reforms is to change the composition rather than the level of labor compensation. The proportion attributable to imputed housing rents has declined, while the percentage formed by wages has increased. Our elasticity estimates suggests that these reforms may have the collateral effect of increasing significantly the labor supply.[25]

These estimates also have policy implications for the evolving discussions of China's welfare system. The estimated sensitivity of work hours to wages demonstrates that urban Chinese labor supplies are quite flexible along the intensive margin (Saez, 2000). The maintained assumption is that urban workers have little discretion, and therefore little flexibility, regarding labor force participation, which is the extensive margin. Therefore, negative income taxes are likely to be more effective than earned income credits. Saez (2000) argues that the former provides minimum incomes to all families at the risk of

[22] We apply the regression-based Hausman test (Davidson and MacKinnon, 1990) for all models; the tests are based on F-statistics. The Hausman test rejects the OLS estimation for all equations presented. The over-identification test does not reject the null hypothesis for all models with one exception, which is identified below.

[23] Regressions that distinguish between non-wage income and wage income of other family members fail to estimate distinct effects of the two in the complete sample and in all subsamples. This is inconsistent with the collective family labor supply model. Browning et al. (1994) and Fortin and LaCroix (1997) reject the unitary family labor supply model in favor of the collective model in Canadian data.

[24] Mokhtari and Gregory (1993) estimate significantly negative uncompensated wage elasticities for labor supply in the Soviet Union. Sharif et al. (1995) suggest that negative uncompensated wage elasticities may also be observed at especially low wage levels. No support for these suggestions appears in these data. The regressions of Tables 4 and 6, with quadratic specifications for own wages, demonstrate that uncompensated labor supply elasticities are insignificant or significantly positive at all wage levels in urban China.

[25] These results cannot provide precise estimates of the expected increases in labor supply because they hold family income, not utility, constant.

discouraging labor force participation. However, this risk appears to be negligible in urban China. Alternatively, earned income credits are essentially negative marginal tax rates on low incomes, and they may encourage workers with higher incomes to reduce labor hours substantially.

Several other explanatory variables also have significant effects on annual work hours. As age increases, people tend to work less. On average, an additional family member is associated with a significant annual increase of approximately six eight-hour work-day equivalents per year. Annual work hours also depend on differences in human capital. More education is associated with significantly fewer annual work hours. The reference group represents the lowest level of educational attainment, i.e., those with less than a lower middle school education. The dummy variables for progressively higher levels of education are associated with negative coefficients of progressively larger magnitude; all are significant at better than the 5% level. These coefficients indicate that the work years of those in the reference group exceed those of workers who graduated from lower middle school by more than one week of eight-hour days on average. This difference increases with the difference in education. For example, the work year of the reference group exceeds that of workers with middle-level professional, technical or vocational degrees by the equivalent of approximately three weeks. Individuals with college or advanced degrees work approximately four and one-half weeks less.

Labor supply also appears to depend on health status. Individuals who receive health care financed by the state provide significantly fewer hours of labor, although the effect is small. An increase in these expenditures equivalent to one standard deviation reduces annual labor supply by only 11.5 hours. In contrast, illness that also entails individual financial expenditures on health care increases annual work hours significantly. An increase in private health care expenditures equivalent to one standard deviation increases annual labor supply by 8.81 hours. Presumably, additional work is necessary to finance these expenditures.[26]

Finally, individuals with private housing work about one and one-half more weeks per year on average than do those who do not have private housing. In addition, peripheral household members supply between four and one-half to five and one-half more weeks of work annually than do household heads, their spouses and children. The remaining variables are insignificant. Labor supply does not depend on the composition or value of assets, if family income is held constant. Work hours do not vary significantly with gender, smoking, Communist party membership or rustification.

The third and fourth columns of Table 4 reproduce the second column separately for men and women, respectively. These estimations demonstrate that the wage and non-labor income effects of the first equation are in the same direction for both genders, but they are markedly stronger for women. The coefficients of both variables are significant for women

[26] Schultz and Tansel (1997) estimate that hours worked in Côte d'Ivoire and Ghana decline significantly with the number of self-reported days disabled. However, they suggest that potential endogeneity may exaggerate the effect, despite their attempts to instrument for days disabled. Our estimates disregard the possibility of endogeneity, in part because we focus on measures of income and wealth, and in part because the possibility of endogeneity between time spent at work and time spent disabled is more troubling than that between the former and explicit expenditures for health care services.

at the 1% level; in contrast, the male wage effect is insignificant and the non-labor income effect is significant at only the 10% level. The female coefficients for both variables are also larger in magnitude than the ones for males. Hence, the implied income and gross wage elasticities are correspondingly larger for women. Our results suggest that women are more responsive to wage and non-wage income changes in urban China, as is the case for the United States (Killingsworth, 1983).

The estimated effects of the remaining variables demonstrate that female labor supply is also more responsive to other influences. The coefficients for age, household head, spouses, children, health care expenditures and private homeownership have the same signs in the male and female equations. However, they are larger in magnitude for women in every case. They are also significant for women in every case, whereas only state healthcare expenditure is significant for men. Some of these differences are substantial. Women who are spouses or children of the household head work nearly 28 fewer days each year than do those who are themselves heads. Women in privately-owned housing have work years that are approximately eleven days longer than those in other forms of tenancy.

In two instances, female responses are significant and of opposite sign to insignificant male effects. Among men, the estimated effects associated with rustification and household status as the parent of the household head are positive but insignificant. These same coefficients are significant and negative for women, indicating reductions of approximately four annual work days for the former and forty days for the latter. Finally, female responses are also greater in magnitude in the one instance in which the male and female coefficients are both significant but have opposite signs. Communist party membership increases male work years by three days but decreases work for women by about four days.

In 1988, household headship status in urban China was almost solely determined by gender. Only 6.13% of women were household heads and only 2.78% of men were spouses of heads. Household organization became noticeably more varied by 1995, when 26.33% of women were household heads and 23.96% of men were spouses of heads (Li and Zax, 2002). Presumably, household heads bear more responsibility for income generation.[27] Therefore, these individuals may exhibit less flexibility in labor supply.

Table 6 stratifies further the equations of Table 4 by headship status. The results are consistent with this hypothesis of inflexibility. The effects of wages and family income are never significant for the labor supply of household heads, regardless of gender.[28] Instead, the strongly significant and positive wage effect in the entire sample is present only in the equations for the labor supplies of male and female non-heads. Similarly, the significant negative effect of family income in the aggregate occurs only among non-heads of either gender. Hence, the absence of these effects in the male equation in Table 4 may be attributable to the preponderance of household heads in the male sample. Similarly, the

[27] Neither the 1988 nor the 1995 surveys define household headship. The survey protocol requires respondents to self-identify as household head or in relation to the head, which is similar to the convention in the US Census. No known source analyzes the determinants of headship designation in these data, even though such an analysis would be interesting. Consequently, we assume that each family identified its own head by choosing the person who best fits that role. The results are consistent with this assumption.

[28] The results for female heads must be interpreted with caution because the test on overidentifying restrictions rejects the null hypothesis for this equation.

Table 6
Labor supply by gender and household responsibility

Variable	Men		Women	
	household head	non-household head	household head	non-household head
Wage (in log term)	−12.81	219.85**	−121.54	210.01**
	(−0.14)	(2.21)	(−1.12)	(3.09)
Non-wage income	−1.76	−6.55*	0.99	−7.22**
	(−0.53)	(−1.89)	(0.22)	(−2.70)
Home assets	0.32	−0.11	0.24	0.0016
	(1.56)	(−0.44)	(0.56)	(0.01)
Age	−4.69	−23.52*	6.36	−27.46**
	(−0.33)	(−1.74)	(0.37)	(−2.66)
Age squared	−0.0039	0.21	−0.14	0.35**
	(−0.03)	(1.28)	(−0.63)	(2.51)
Years worked	2.28	2.58	4.92*	−1.57
	(1.12)	(0.89)	(1.93)	(−0.65)
Party affiliation	35.31**	6.80	−20.49	−23.93
	(2.47)	(0.28)	(−0.86)	(−1.06)
Family size	19.48	65.68**	24.81	59.88**
	(1.06)	(3.37)	(1.45)	(4.24)
Home ownership, public	−44.42	49.75	51.72	42.65
	(−0.90)	(0.76)	(0.61)	(0.76)
Home ownership, private	−13.93	111.39	55.72	105.18*
	(−0.27)	(1.61)	(0.65)	(1.82)
Healthcare cost, individual	0.042*	−0.018	0.050*	0.032
	(1.92)	(−0.48)	(1.78)	(1.32)
Healthcare cost, state	−0.0085*	−0.012	0.00059	−0.017**
	(−1.75)	(−0.97)	(0.08)	(−2.65)
Rusticated youth	−9.14	12.64	0.98	−53.005**
	(−0.45)	(0.55)	(0.06)	(−2.49)
Cigarettes per day	0.11	1.79*	−0.22	0.70
	(0.15)	(1.71)	(−0.08)	(0.25)
Education level, university	−164.13**	−167.96*	−107.66	−178.96**
	(−3.29)	(−2.04)	(−1.36)	(−3.17)
professional	−168.01**	−98.24	−71.39	−183.32**
	(−4.50)	(−1.36)	(−1.04)	(−3.92)
technical	−150.91**	−85.07	−50.85	−141.27**
	(−3.66)	(−1.17)	(−0.78)	(−2.97)
upper middle	−128.19**	−44.54	−67.13	−23.25
	(−3.55)	(−0.65)	(−1.10)	(−0.58)
lower middle	−103.02**	11.76	−45.44	−12.64
	(−3.11)	(0.18)	(−0.80)	(−0.33)

(*continued on next page*)

large majority of non-heads in the female sample is responsible for the presence of these effects in the female equation in Table 4.

In other words, individual labor supply responses to changes in wage and non-labor income are clearly associated with family responsibilities. From Table 6, the estimated elasticities of labor supply with respect to family income, gross and compensated wages

Table 6 (*Continued*)

Variable	Men		Women	
	household head	non-household head	household head	non-household head
Observations	3167	2407	1627	3359
F-statistic	$(19, 3147) = 3.87$	$(19, 2387) = 3.35$	$(19, 1607) = 2.21$	$(19, 3339) = 4.34$
Prob > F	0.00	0.00	0.0023	0.00
Overidentifying restriction test (F-stat)	$(6, 3141) = 0.36$	$(6, 2381) = 0.57$	$(6, 1601) = 3.94$	$(6, 3333) = 0.43$
Prob > F	0.90	0.75	0.0006	0.86
Hausman test (F-stat)	$(2, 3145) = 21.40$	$(2, 2385) = 30.19$	$(2, 1605) = 10.80$	$(2, 3337) = 48.38$
Prob > F	0.00	0.00	0.00	0.00
Elasticity, calculated				
Income	−0.009	−0.040	0.005	−0.044
Gross wage	0.006	0.100	−0.057	0.096
Compensated wage	6.14	18.64	−3.066	18.51

Notes. 1 The constant term is not reported. All standard errors are heteroskedasticity robust, and t-statistics are in parentheses. 2. The elasticity is calculated based on the relevant sample averages of working hours, non-wage family income, and wages.

* Significance at the 10% level.

** Significance at the 5% level.

are essentially identical for female and male non-head household members. The respective coefficients are also similar for female and male heads in terms of significance, although the magnitudes are quite different.[29] In particular, compensated wage increases yield much larger increases in labor supply for male and female non-heads than for heads. For the former, the increase in the price of leisure associated with increasing wages dominates the parallel increase in income. For household heads, the income and substitution effects of wage increases are nearly offsetting. This result suggests that the marginal value of leisure may be higher for workers having more responsibilities for family income. The fact that annual average work hours are fewer for household heads may be due to more opportunities for joint consumption with other family members. Age and the number of family members also affect labor supply only for non-heads. The effect of the remaining variables is less consistent. For example, male household heads demonstrate significantly reduced labor supplies at all educational levels above the reference, but these reductions are present only for the three highest educational categories among female non-heads and only for university graduates among male non-heads. Labor supplies do not vary significantly with education among female household heads.

In summary, our results demonstrate, first, that OLS estimations of labor supply equations in urban China are seriously misleading. They indicate incorrect signs for both wages and non-wage income. Tests indicate that they are contaminated with endogeneity bias. In contrast, our IV estimations rely on instruments that appear to be valid. These estimations identify positive compensated wage effects and negative income effects on work hours. Income and gross wage effects are very small. Nevertheless, estimated

[29] Table 5 reports results for pooling male and female household heads and non-heads.

compensated wage effects are quite large. However, wage and income effects on labor supply depend heavily on household responsibilities. The labor supplies of household heads, who presumably bear most of these responsibilities, are relatively insensitive to changes in wages and non-wage income. The apparent effects of these changes in equations pooled over all sample members are largely attributable to the labor supplies of non-household heads.

6. Conclusions

Estimating labor supply elasticities is often a frustrating task. In addition to econometric difficulties, workers often have only restricted opportunities to adjust working hours in response to changes in their circumstances. Much of the flexibility in labor supply occurs through the choice of jobs, occupations, and ownership sectors, rather than through adjustments in hours on a given job. Hence we might not expect to find significant labor supply responses in the incompletely reformed urban Chinese labor market of 1995.

Using micro-level data, we investigate the determinants of labor supply in urban China in 1995, placing particular emphasis on the effects of wages and non-labor incomes. Instrumental variables estimation reveals substantial negative income effects and positive wage effects. Our results indicate that consumption of leisure declines when its price increases without compensation and that leisure is a normal good. Moreover, they are consistent with economic theory. Compensated wage effects on labor supply are strongly positive. At the low levels of income typical in urban China, compensated increases in the cost of leisure encourage substantial substitution towards the consumption of purchasable commodities. Empirically, these effects depend heavily on family responsibilities. The labor supplies of household heads of either gender are not sensitive to wages and family income changes. All significant responses pertain to non-heads only. In the aggregate, female labor supplies appear to be more responsive, but this contrast is attributable to the relative scarcity of household heads among women.

Our findings have important policy implications. During the course of economic transition, reforms inevitably change wages or non-wage income. Income and wage elasticities provide a basis for evaluating the labor supply consequences of these policies. Large compensated wage elasticities suggest that labor supplies of non-household heads will increase substantially if Chinese reforms continue to substitute wages for entitlements in state-owned enterprises. Small uncompensated wage elasticities indicate that increases in income or social security taxes are likely to be born by workers rather than passed on to employers. Finally, labor supply responses will become more important as continuing reforms remove some of the rigidities that remained in the urban Chinese labor market in 1995.

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References

Apps, Patricia F., Rees, Ray, 1997. Collective labor supply and household production. J. Polit. Economy 105 (1), 178–190.

Basmann, Robert L., 1960. On finite sample distributions of generalized classical linear identifiability test statistics. J. Amer. Statist. Assoc. 55 (292), 659.

Blundell, Richard, MaCurdy, Thomas, 1999. Labor supply: a review of approaches. In: Ashenfelter, Orley C., Card, David (Eds.), Handbook of Labor Economics, Vol. 3A. North-Holland, New York, pp. 1559–1695.

Bound, John, Brown, Charles, Duncan, Greg J., Rodgers, Willard J., 1990. Measurement error in cross-sectional and longitudinal labor market surveys: validation study evidence. In: Hartog, Joop, Ridder, Geert, Theeuwes, Jules (Eds.), Panel Data and Labor Market Studies. North-Holland, New York, pp. 1–19.

Bound, John, Krueger, Alan B., 1991. The extent of measurement error in longitudinal earnings data: do two wrongs make a right? J. Lab. Econ. 9 (1), 1–24.

Browning, Martin, Bourguignon, Francois, Chiappori, Pierre-Andre, Lechene, Valerie, 1994. Income and outcomes: a structural model of intrahousehold allocation. J. Polit. Economy 102 (6), 1067–1096.

Byron, Raymond P., Manaloto, Evelyn Q., 1990. Returns to education in China. Econ. Devel. Cult. Change 38 (4), 783–796.

Burkett, John P., Putterman, Louis, 1993. The case of Dahe commune. Economica 60 (240), 381–396.

Chiappori, Pierre-Andre, 1988. Rational household labor supply. Econometrica 56 (1), 63–90.

Chiappori, Pierre-Andre, 1992. Collective labor supply and welfare. J. Polit. Economy 100 (3), 437–467.

Chiappori, Pierre-Andre, 1997. Introducing household production in collective models of labor supply. J. Polit. Economy 105 (1), 191–209.

China Statistical Publishing House, 1997. China Labor Statistical Yearbook. Beijing, China.

Conway, Karen S., Kimmel, Jean, 1998. Male labor supply estimates and the decision to moonlight. Lab. Econ. 5 (2), 135–166.

DaVanzo, Julie S., DeTray, Dennis N., Greenberg, David H., 1976. The sensitivity of male labor supply estimates to choice of assumptions. Rev. Econ. Statist. 58, 313–325.

Davidson, Russell, MacKinnon, James G., 1990. Specification tests based on artificial regressions. J. Amer. Statist. Assoc. 85, 220–227.

Davidson, Russell, MacKinnon, James G., 1993. Estimation and Inference in Econometrics, Chapter 7. Oxford Univ. Press, New York/Oxford.

Dickens, William T., Lundberg, Shelly J., 1993. Hours restrictions and labor supply. Int. Econ. Rev. 34 (1), 169–192.

Dong, Xiao-yuan, Dow, Gregory K., 1993. Monitoring costs in Chinese agricultural teams. J. Polit. Economy 101 (3), 539–553.

Fortin, Bernard, LaCroix, Guy, 1997. A test of the unitary and collective models of household labor supply. Econ. J. 107 (443), 933–955.

Hansen, Lars Peter, 1982. Large sample properties of generalized method of moments estimators. Econometrica 50 (4), 1029–1054.

Kahn, Shulamit, Lang, Kevin, 1992. Constraints on the choice of work hours: agency versus specific capital. J. Human Res. 27 (4), 661–678.

Killingsworth, Mark R., 1983. Labor Supply. Cambridge Univ. Press, New York.

Kinal, Terrence W., 1980. The existence of moments of K-class estimations. Econometrica 48 (1), 241–249.

Knight, John, Song, Lina, 1991. The determinants of urban income inequality in China. Oxford Bull. Econ. Statist. 53 (2), 123–154.

Korenman, Sanders, Neumark, David, 1991. Does marriage really make men more productive? J. Human Res. 26 (2), 282–307.

Lazear, Edward P., 1981. Agency, earnings profiles, productivity, and hours restrictions. Amer. Econ. Rev. 71 (4), 71–85.

Li, Haizheng, 2000. Economic efficiency and social insurance reforms in China. Contemporary Econ. Pol. 18 (2), 194–204.

Li, Haizheng, Zahniser, Steven, 2002. The determinants of China's temporary rural–urban migration. Urban Stud. 39 (12), 2219–2235.

Li, Haizheng, 2003. Economic transition and returns to education in China. Econ. Educ. Rev. 22 (3), 317–328.

Li, Haizheng, Zax, Jeffrey S., 2002. Economic transition and the labor market in China. Working paper. School of Economics, Georgia Institute of Technology.

Liang, Zai, White, Michael J., 1997. Market transition, government policies, and interprovincial migration in China: 1983–1988. Econ. Devel. Cult. Change 45 (2), 321–339.

Liu, Minquan, 1991. Intersectoral labor allocation on China's communes: a temporal priority analysis. J. Compar. Econ. 15 (4), 602–626.

Liu, Zhiqiang, 1998. Earnings, education, and economic reforms in urban China. Econ. Devel. Cult. Change 46 (4), 697–725.

Makunnas, Seija, Pudney, Stephen, 1990. A model of female labour supply in the presence of hours restrictions. J. Public Econ. 41 (2), 183–210.

Mokhtari, Manouchehr, Gregory, Paul R., 1993. Backward bends, quantity constraints, and Soviet labor supply: evidence from the Soviet interview project. Int. Econ. Rev. 34 (1), 221–242.

Mroz, Thomas A., 1987. The sensitivity of an empirical model of married women's hours of work to economic and statistical assumptions. Econometrica 55 (4), 765–799.

Nakamura, Alice, Nakamura, Masao, 1981. A comparison of the labor force behavior of married women in the United States and Canada, with special attention to the impact of income taxes. Econometrica 49 (2), 451–489.

Pencavel, John, 1986. Labor supply of men: a survey. In: Ashenfelter, Orley C., Layard, Richard (Eds.), Handbook of Labor Economics, Vol. 1. North-Holland, New York, pp. 103–204.

Putterman, Louis, 1990. Effort, productivity, and incentives in a 1970s Chinese people's commune. J. Compar. Econ. 14 (1), 88–104.

Rodgers, Willard L., Brown, Charles, Duncan, Greg J., 1993. Errors in survey reports of earnings, hours worked, and hourly wages. J. Amer. Statist. Assoc. 88 (424), 1208–1218.

Saez, Emmanuel, 2000. Optimal income transfer programs: intensive versus extensive labor supply responses. Working paper No. 7708. National Bureau of Economic Research.

Sahn, David E., Alderman, Harold, 1996. The effect of food subsidies on labor supply in Sri Lanka. Econ. Devel. Cult. Change 45 (1), 125–145.

Schultz, Paul T., Tansel, Aysit, 1997. Wage and labor supply effects of illness in Côte D'Ivoire and Ghana: instrumental variable estimates for days disabled. J. Devel. Econ. 53 (2), 251–286.

Sharif, Mohammed, Suzawa, Gilbert S., Miller, Carole F., 1995. Forward-falling labor supply implications for wage rigidity, unemployment, and plan failure. J. Econ. Devel. 20 (1), 91–112.

Tummers, Martijn P., Woittiez, Isolde, 1991. A simultaneous wage and labor supply model with hours restrictions. J. Human Res. 26 (3), 393–423.

Xie, Xiaodi, 1997. Children and female labour supply behavior. Appl. Econ. 29 (10), 1303–1310.

Yang, Dennis Tao, Zhou, Hao, 1999. Rural–urban disparity and sectoral labour allocation in China. J. Devel. Stud. 35 (3), 105–133.

Zax, Jeffrey S., 1995. Human capital in a workers' paradise. Working paper. Univ. of Colorado at Boulder.

Zax, Jeffrey S. Housing Reform in Urban China. In: Nicholas Hope (Ed.), How Far Across the River? Chinese Policy Reform at the Millennium. Stanford Univ. Press, Palo Alto. In press.

Zhao, Yaohui, 1999. Leaving the countryside: rural-to-urban migration decisions in China. Amer. Econ. Rev.-Papers and Proceedings 89 (2), 281–286.