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Examining the Timing of Women's Retirement in Urban China: A Discrete Time Hazard Rate Approach

Denise Hare

Economics Department, Reed College, 3203 SE Woodstock Blvd., Portland, OR 97202

Abstract

In spite of phenomenal wage growth in recent decades, female labor force participation in urban China has experienced a notable and puzzling decline, particularly so for women aged 40 and above. Concerns about women's status motivate an inquiry into the reasons behind the sharp drop, and consideration of how this decline may bear on their welfare. This paper estimates a probit model of labor force participation and a discrete time hazard rate model of retirement and non-retirement labor force withdrawals among urban married women aged 40 to 51. The analysis uses data from a multi-province panel data set covering the interval from 1991 to 2011. The results inform policy discussions about pension system reform in the face of population aging and declining labor supply. In addition, recommendations are offered for policy measures that would increase women's labor force attachment, even as they approach their retirement horizon.

Keywords

China; gender; retirement; urban; women; labor force participation

I. Introduction

One of the most striking recent labor market developments in urban China is a marked decrease in women's labor force participation. This is in the face of astonishingly high and persistent economic growth in the post reform era, which drives steady increases in wage rates. The drop off in participation is particularly steep for women aged 40 and above, with many reporting that they are retired or otherwise appearing permanently withdrawn from the labor force. While a modest body of scholarship examines women's labor force participation, employment, and wages in urban China more generally (He and Zhu, 2016; Chi and Li, 2014; Dong and Jia, 2013, and Maurer-Fazio, *et al.*, 2011 are some recent examples), very little attention has been paid to the specific circumstances of women over 40, and their unusually high incidence of withdrawal, although Giles, *et al.* (2011) document that participation among older women is declining over time.

China's urban workers, admittedly, face relatively low official retirement ages; for many women the customary age is 50, which also represents the start of pension eligibility. However, the current pension system is plagued with fiscal challenges and uncertainties

about its future direction. Being neither uniform nor centrally managed (Cai and Cheng, 2014), coverage is far from complete. It is reported that only 55.9 percent of the urban employed population participated in a pension program in 2010 (Giles, *et al.*, 2013, p. 8), and even among those who are covered, the rules and formulas that determine benefit levels can vary widely by region (Cai and Cheng, 2014; Giles, *et al.*, 2013). It is well-known that programs in many areas are plagued by funding shortfalls, leaving benefit receipt uncertain, and population aging will only exacerbate the threats of insolvency (Cai and Cheng, 2014; Pozen, 2013; *The Economist*, 2014).

Variability in benefits, difficulty in coordinating benefits for workers with multiple employers (and particularly with work histories in more than one region), and uncertainty about the sufficiency of local insurance pools all contribute to a sense of distrust, culminating in a general reluctance to join the system or to rely on it for old-age support (*The Economist*, 2014; Pozen, 2013). Though participation in the basic urban pension insurance program is mandated for any wage-earning employee of a firm with 8 or more workers (Giles, *et al.*, 2013), it is important to note that both compliance and enforcement cannot be taken for granted. Historically, public sector employees have enjoyed very favorable contribution terms (Roberts, 2013), and therefore participation is understandably high in that sector. Pension coverage is found to be lowest among workers employed in the private sector and among the self-employed (Giles, *et al.*, 2013). Exacerbated by China's aging population, the precarious balance of the system relies on a capacity for facilitating longer working years and later retirements among urban workers (Shi, *et al.*, 2015). In fact, China's aging population raises further concerns about a declining labor supply going beyond just its implications for the pension system (Du and Yang, 2014; Cai and Lu, 2013). Though increases to retirement and pension eligibility ages are expected, no concrete plans for changes have been released (Chen, 2017).

The results of this paper contribute towards the pension system reform discussion by demonstrating how rules concerning retirement age may bear on women's labor force withdrawals as they approach their retirement horizon. For instance, official retirement ages differ by occupation, leading to variation across women in the age benchmark that they face. This variation enables me to examine how retirement age affects withdrawal decisions. Women who are not covered by the current pension system, however, may be less responsive to age benchmarks because they will not receive a pension benefit of any amount no matter how long they work. Although a positive correlation between pension eligibility and retirements among individuals in urban China has been established (Giles, *et al.*, 2015) until now, there have been no studies which examine the timing decisions of women's (or men's) retirements, and how they may be affected by the parameters of pension system policy. My study fills this gap.

Beyond consideration of pension system reform, several additional policy issues arise in connection with the patterns of women's labor force withdrawals examined here. Alongside the dismantling of the social safety net, women have shouldered a disproportionate share of care work, for both young and elderly dependents, which may interfere with their opportunities for paid employment (Dong and An, 2012). Economic liberalization along with state sector restructuring have exerted wide ranging influences on labor markets, many

of which have been reported to be discouraging or even discriminatory towards women's employment (Dong and Pandey, 2012; Du and Dong, 2009). Changing wage structures, reflecting global trends, increasingly disincentivize labor force participation among the less skilled and less educated (Hare 2016; Ge and Yang, 2014).

This paper seeks to explain the observed pattern of labor force withdrawals among urban married women, between ages 40 and 51, using micro-level data from eight waves of the China Health and Nutrition Survey (CHNS) administered between 1991 and 2011. Though certain aspects of the survey design dictate the sample cutoff at age 51, the sharpest decline in labor force participation occurs within this interval, and the number of women continuing to participate beyond age 51 is very small. The analysis will identify factors that account for women's labor force withdrawal as they age through this interval. The panel dimension of the data, coupled with its generous (better than 20 year) timespan, facilitates examination of how changing economic conditions have influenced withdrawal patterns. I proceed to review the relevant literature, followed by presentation of the data, methods, and results, before finally concluding with a discussion of policy implications.

II. Literature review

Retirement outcomes among China's urban women are shaped by many factors. For instance, the eligibility requirements, benefit levels, and sectoral coverage of the current pension system all undoubtedly exert influence on the timing of labor force withdrawals. China's changing economic landscape also plays a role as state enterprise retrenchment coupled with increased global integration have led to shifting patterns of labor demand which may have distinctly gendered impacts. At the same time, demographic characteristics contribute to the opportunity cost of time and the formation of women's reservation wages, particularly with respect to the care needs of the youngest and eldest family members.

As would be expected, own pension eligibility is positively correlated with women's retirement, though somewhat surprisingly, the pension eligibility of husbands fails to exert a significant effect (Giles, *et al.*, 2011). The lack of influence from husbands' pensions is consistent with findings drawn from the analysis of U.S. data. Gustman and Steinmeier (2000) as well as Coile (2004) note an asymmetry in spousal effects, with elements of the wife's decision having influence on the husband's retirement, but not vice versa. The evidence supports the notion of coordination in retirement timing, which is attributed to leisure complementarities or to a correlation in preferences (Gustman and Steinmeier, 2000). Turning back to the case of China, the employment of husbands is shown to increase, by about 16 percentage points, the likelihood of working among women age 45 and above (Giles, *et al.*, 2015), consistent with the hypothesis of retirement timing coordination.

The realignment of state and collective owned enterprises starting from the mid to late 1990s gave rise to changing patterns of urban employment and labor force attachment. Retrenchment led to the discharge of literally millions of urban workers through the late 1990s and early 2000s (Dong and Putterman, 2003; Dong and Xu, 2009; Giles, *et al.* 2006a; Solinger, 2002). It has been reported that both women and older workers were particularly hard hit by enterprise lay-offs and closures (Dong and Pandey, 2012; Giles, *et al.*, 2006a).

Women also experienced more challenging prospects for re-employment (Du and Dong, 2009), with older age being an especially strong deterrent to women's re-employment (Giles, *et al.*, 2006b). Hence, the advent of the new millennium gave rise to a much less certain employment scenario than many urban workers had experienced previously, and women, especially older women, were among the most vulnerable.

Enterprise restructuring was accompanied by structural change, coincident with China's accession to the WTO at the end of the year 2000. As resource allocation decisions have become increasingly subject to market price signals, the factors that determine labor wage payments have experienced realignment. Several recent papers have noted both a widening of the wage distribution along with increases in the returns to skill and education (Appleton, *et al.*, 2014; Ge and Yang, 2014; Meng, *et al.*, 2013; Wang, 2013). Other studies (Connelly, *et al.*, 2014; Giles, *et al.*, 2015) provide corroborating evidence that, in the face of market churning, higher schooling achievement, particularly upper secondary completion and beyond, increases labor force participation among older women in urban China.

Even as their employment situation was in flux, residents of urban China also were subject to changes in the social safety net and the public provision of social services, which some argue have fundamentally changed the obligations they face with respect to the care needs of their young and elderly family members (Du and Dong, 2013). Connelly, *et al.*, (2014) and Giles, *et al.*, (2015) report that the presence of grandchildren yields a negative impact on labor force participation among older women. Chen, *et al.*, (2011) confirm the important role of grandparents as child caregivers. They report that hours of childcare provided by coresident grandmothers actually exceeds that of mothers, except for the case of children under age 1. With respect to non-coresident grandchildren, paternal grandparents account for the 27 percent of the childcare provided by non-household members, and maternal grandparents account for 13 percent. Exploiting variation in the age of pension eligibility, Feng and Zhang (2017) show that retirement confers a 29 percentage point increase in urban women's likelihoods of providing grandchild care.

While the net effect of coresidence with elders is shown to have a positive effect on urban women's labor force participation (Maurer-Fazio, *et al.*, 2011), elder care needs are negatively correlated with their employment status and hours of work (Liu, *et al.*, 2010). Dong and An (2012), analyzing time use data from 2008, confirm that women are more susceptible than men to picking up the responsibility for care obligations and other housework tasks, and, furthermore, that the number of hours devoted to this class of activities is increasing in age.

Additional support for the connection between care obligations and retirement decisions can be found in the broader literature. Drawing on representative household level data from Germany, Meng (2011) shows that when an employed women aged 58 or above is engaged in informal care giving, her hazard rate of retirement is raised by a striking 74 percent in the subsequent year. Analyzing U.S. data, Lumsdaine and Vermeer (2015), Rupert and Zanella (2016), and Asquith (2016) all demonstrate negative correlations between the birth of a grandchild and grandparent labor supply. Asquith (2016) reports that the likelihood of women's labor force withdrawal increases by as much as 13.8 percent in response to the

birth of each grandchild. In contrast, when women provide financial support towards younger or elder family members, the sign between giving support and leaving the labor force reverses, since income from paid work facilitates the resource transfer. Henretta and O'Rand (1980), also using U.S. data, find that supporting either a parent or child reduces the probability of stopping work by as much as 6.2 percentage points for women under age 58 and by as much as 13.9 percentage points for women age 58 and over.

The body of work reviewed here raises several testable hypotheses about the determinants of women's withdrawal from the labor force as they approach their retirement horizon. In the next section I will proceed to identify determinants of labor force withdrawal, using both static and dynamic approaches. From a policy perspective, a critical issue revolves around learning what obstacles and incentives are at play in producing the observed pattern of labor force withdrawals. I now turn to examination of the data.

III. Data and methods

The China Health and Nutrition Survey (CHNS), administered by the Carolina Population Center (University of North Carolina) in collaboration with the National Institute of Nutrition and Food Safety (Chinese Center for Disease Control and Prevention), is the source of the data analyzed here. Household respondents were selected through a multi-stage random cluster process, and although both rural and urban households are included, I utilize only the urban portion of the sample. Though not designed to be nationally representative, steps were taken to ensure a wide range of socioeconomic circumstances among selected households (Popkin, *et al.*, 2010). Cities and counties were chosen randomly following their stratification into low-, middle-, and high-income groups. At the next stage, communities (neighborhoods in the case of the urban sample) were drawn randomly, and finally, 20 households from each community were chosen, at random, for an interview.

The survey was first implemented in 1989, though because key aspects of the 1989 wave are incompatible with subsequent years, my analysis begins with the 1991 wave and draws from 7 subsequent waves¹ through the most recent conducted in 2011. In each wave, between 3,000 and 5,000 households were sampled. The survey was designed as a panel study, but due to attrition, new respondents were added, starting in 1997 and in subsequent wave years, to replace those that were lost. Attrition arises for several reasons here, including the inability to reach people who are far away from home (for work, travel, or school) at the appointed survey time; the incidence of natural disasters which have impeded access to some survey sites in certain wave years; and, finally, difficulty in finding respondents subsequent to housing relocation associated with some urban redevelopment projects (Popkin, *et al.*, 2010). Over the years, nine different provinces (Liaoning, Heilongjiang, Jiangsu, Shandong, Henan, Hubei, Hunan, Guangxi, and Guizhou) are represented, and in the most recent 2011 wave, the cities of Beijing, Shanghai, and Chongqing were added.

Throughout the paper I focus solely on women reporting to be married, thereby eliminating the never married, divorced, widowed, or separated. The large majority of women in the age

¹Survey wave years included in my analysis are 1991, 1993, 1997, 2000, 2004, 2006, 2009, and 2011.

group I examine are married, and because spousal and familial characteristics are key elements of my estimation exercises, the analyses do not lend themselves well to unmarried women. There is a total of 10,550 observations on urban married women between the ages of 25 and 64, and all the analyses conducted in this paper will be performed on various subsamples drawn from this set of observations. Online Appendix Table A describes all of these subsamples and the number of observations contained in each, noting the figures and tables for which each subsample is utilized.

Labor force participation rates (calculated as three-year moving averages) for urban married women, aged 25 through 64, are plotted in Figure 1. To simplify the presentation, only three wave years of data are utilized in this figure: the endpoint wave years of 1991 and 2011, along with the 2000 wave year from the middle of the sample. A distinct age profile is apparent across all three years of the data, with rates of participation highest among women up until age 40, when participation begins to fall—very steeply after age 45, followed by some flattening beyond age 55. The number of observations in each cell gets smaller with increasing age, and so observed fluctuations may be less representative of population trends in the very upper age range.

In order to focus on women's decisions as they approach their retirement horizon, the remaining analysis will be restricted to urban married women from age 40 to 51. The upper age limit might have been set slightly higher, but a key survey module was executed only for women below age 52, and as a result some variables that are necessary for my analysis cannot be included beyond age 51. However, as shown by the graph in Figure 1, a notably large portion of urban married women experience labor force withdrawal at some point within this interval. Moreover, women in blue-collar positions face a mandatory retirement age of 50, though white-collar women may retire at age 55. Narrow exceptions to these rules allowing for a retirement age of 60 is reserved for women in select positions such as university professor (Giles, *et al.*, 2011).

There is a total of 3,431 individual-year observations on married urban women age 40 to 51, representing 1,813 unique women. Among these, labor force status is missing for 9 individual-year observations and so these are omitted completely from the analysis. In the remaining set of 3,422 observations, 1,153 are not in the labor force, 2,193 are employed, and 76 are unemployed. Online Appendix Figure A illustrates the reasons given for non-participation in the labor force when restricting to the 1,153 labor force non-participants between ages 40 and 51. The portion reporting that they are retired is slightly more than one-third of the sample, while nearly 47 percent report that they are doing housework. Unknown or other reasons account for the remainder.

Switching our focus to the 2,193 employed observations, Online Appendix Figure B illustrates the composition of employment sectors among women aged 40 to 51. The largest portion, 50 percent, reports government employment. This represents what we conventionally call the public sector, containing governing bodies (legislative, judicial, and executive branches) and administrative departments and bureaus, in addition to service entities such as public universities, hospitals, and research institutes. The second largest category is the self-employed group at 14 percent. State- and collectively-owned enterprises

each account for about 13 percent and 7 percent, respectively. State-owned enterprises are engaged in commercial or production activity with assets held under central government control, while collectively-owned enterprise assets are typically held at a municipal level. A little more than 11 percent of women are employed in privately held or foreign-owned enterprises, and the sector of employment of the remainder, 4 percent, is unknown. These figures are reasonably consistent with published national statistics. For instance, selecting the year 2001, which falls midway through my 1991–2011 study interval, state-owned units accounted for 65.98 percent of urban female employment (National Bureau of Statistics, 2002, Table 1–10). The national categories are not broken down as finely as those in the CHNS survey data, rather state-owned represents an amalgamation of the government category joined together with the state-owned enterprise category. Therefore, the 65.98 percent national measure should be compared to 49.99 plus 13.13, or 63.12 percent from the CHNS subsample. Meanwhile, 12.07 percent of urban females are employed in collectively-owned units, nationally, compared to 6.66 percent in my CHNS data sample, while 21.95 percent are employed in the remaining “other” ownership category, nationally, compared to 30.24 percent in my CHNS data sample. By 2011, the contribution of state-owned units to urban female employment, nationally, had fallen to a 48.25 percent share (National Bureau of Statistics, 2012, Table 1–14) as a result of state sector restructuring.

Online Appendix Figure C displays the occupational status of women in my CHNS subsample, and the largest portion here, 46 percent, reports being engaged as a production or service worker. Women in the second largest group, at 21 percent, hold professional or technical positions, while 10 percent are administrators or managers. Office staff makes up 14 percent of the group, and occupation is unknown for the remaining 8 percent.

Further analysis of the data will employ two distinct estimation exercises. The first uses a bivariate probit regression to analyze the labor force participation status of urban married women aged 40 to 51 in order to learn how women who are participating in the labor force differ from those who do not. This analysis draws on the pool of 3,431 urban married women between age 40 and 51. Among these, labor force status is not observed for 9 observations and there is missing data on one or more independent variables for another 286. Therefore, my estimation sample contains 3,136 individual-year observations representing 1,701 unique women. Online Appendix Table B contains the sample means and standard deviations of the dependent and independent variables.

In the second exercise, a discrete time hazard model is utilized, which estimates the conditional probability that a failure event occurs at time t , given that it has not occurred prior to time t . The event to be examined is labor force withdrawal. The measurement of withdrawal status at time t is based on the observation of the woman’s labor force participation in the subsequent wave year. A woman is assigned “withdraw” if she withdraws during the interval leading to the next wave’s observation (i.e., if she is no longer participating at the next wave). She is recorded as “does not withdraw” if she is still participating in the labor force at the next wave. Due to its many desirable features, hazard rate estimation is widely used for examining the length of spells preceding transitions in and out of employment or labor force participation. The advantage of the dynamic hazard rate approach relative to the static probit analysis lies in its ability to identify the conditions

associated with withdrawal timing, recognizing that a decision to stay maintains the possibility of later withdrawal (Lumsdaine and Vermeer, 2015).

The discrete (versus continuous) time model is appropriate, given the interval nature of the data, because it allows the covariates to update, as needed, at each observed sample wave, while the logit specification offers a very flexible functional form allowing for the possibility of non-proportional hazards (Jenkins, 1995). The hazard rate function is given by:

$$h(t, \mathbf{X}_{i,j,t}) = [\exp(\alpha_t + \mathbf{X}_{i,j,t}\beta)] / (1 + \exp(\alpha_t + \mathbf{X}_{i,j,t}\beta))$$

where the vector α_t is comprised of a set of indicator variables representing the length of the spell (number of survey waves) for which the woman has been observed up until year t , while the vector $\mathbf{X}_{i,j,t}$ consists of characteristics of individual i in province j at year t . Contained in $\mathbf{X}$ are variables measuring individual and household characteristics that reflect both the supply (reservation wage) and demand (offered wage) labor market conditions that bear on women's decisions about whether to continue to participate in the labor force or to withdraw. Also included are sets of province (j) and year (t) fixed effects. Sample exit takes place when the woman either withdraws or ceases to be observed in the sample. Women who exit the sample before withdrawing contribute to the estimation as censored data whereas women who are observed to withdraw contribute completed duration data. Finally, the possibility of unobserved heterogeneity is allowed through the estimation of a random effect for each unique woman in the sample.

Constructing the estimation sample for the hazard rate estimation exercise, I begin with a universe of 3,422 individual-year observations on urban married women of age 40 to 51 for whom labor force status is observed. Among these there are 1,813 unique women. From this universe, 2,344 individual-year observations are lost through the series of exercises undertaken to define the hazard analysis sample. (Because the failure event is coded from the subsequent wave, a substantial portion of the sample reduction arises because the last recorded wave on each individual cannot enter the estimation sample.) This yields a set of 1,078 individual-year observations. However, within this group, only two women report an episode of unemployment, and since we cannot suitably control for this factor with so few observations, they are dropped also (both women experience withdrawal subsequent to being unemployed). The final estimation sample contains 1,076 individual-year observations on 645 unique women. Online Appendix Table B contains the means and standard deviations of the dependent and independent variables for each of the estimation model samples. By construction, the hazard analysis sample is considerably smaller than labor force participation probit model sample. Worries about selection bias are not severe, however, because although not identical, the variable means are reasonably similar across the two. One notable difference is that the hazard rate estimation sample is younger, and this arises because construction of the sample forces exit upon the incidence of a labor force withdrawal. Hence, observations on withdrawn women are dropped, and these tend to be older women.

The incidence of the hazard model failure event, withdrawal, is illustrated in Figure 2, where age is on the horizontal axis, while the vertical axis on the left side measures the percentage of those women observed to have withdrawn by the subsequent survey wave. The right side vertical axis measures the number of observations on women at each age (from 40 to 51). The number of observations, of course, declines as women experience the withdrawal event. The decrease is non-monotonic due to the interval nature of the data collection. For example, our sample contains only 108 women of age 42, while the number of women observed at age 43 is 135. Because data were collected at 2 to 4 year intervals, these two groups do not contain any of the same women. A 42-year old woman is not observed again until she is 44, 45, or 46 years of age.

Turning to the rate of withdrawals, we observe a non-monotonic increase for women through most of their forties. There is a small bump in the early forties. Withdrawals begin to rise again immediately after age 45, reaching a peak at 48 (where the withdrawal rate reaches 50 percent). This pattern is consistent with a general retirement age of 50 years for most women (recall that the actual withdrawal takes place in the interval leading to the next wave year).

Many, but not all, of the sample women report that they are retired coincident with their labor force participation withdrawal. Of the 238 in our hazard analysis estimation sample for whom a withdrawal event occurs, 154 of them respond with the answer of “retired” as the reason for not working, while 84 give “doing housework” or “other/unknown” as the response to this question. Breaking down the withdrawal event into two separate components—self-reported retirements and other non-retirement withdrawals—allows for further inspection of withdrawal trends. Non-retirement withdrawals (defined as those who respond “doing housework” or “other/unknown” as their reason for not working) are illustrated in Figure 3. Though we have redefined the failure event, the at-risk sample remains unchanged and so the number of observations, at each age, is identical to that presented in Figure 2. Non-retirement withdrawals demonstrate little sensitivity to age, although the highest rates of non-retirement withdrawals occur among women in their early 40s. Nonetheless, the peak, at age 43, is only about 13 percent. Retirement withdrawals (Figure 4), in contrast, demonstrate a much stronger correlation with age. The sharpest rise occurs subsequent to age 45, with better than 42 percent of women at age 48 being retired by their next wave observation. Further exploration of the patterns encountered here is the subject of the next section.

IV. Estimation results

Labor force participation modeling utilizes a reduced form approach, where covariates proxy for factors on both the supply (reservation wage) and demand (offered wage) sides of the labor market. Human capital is expected to influence participation through offered wages and is measured by a set of binary variables that represent incremental levels of schooling completion. The final education category is constructed to include both upper secondary vocational or technical training as well as any sort of post-secondary schooling. While these may represent somewhat different types of skill acquisition, I found their coefficient estimates to be statistically indistinguishable from one another, and in later stages of the

analysis, small cell size makes it impossible to estimate separate coefficients, therefore these categories are combined into one.

Non-labor income is measured at the household level, deflated to 1988 constant *yuan* values, and represents the sum of income from all non-labor sources such as assets, transfers, and gifts. It is expected to influence participation through an income effect on the reservation wage. The variable measuring difference in age between husband and wife tests for evidence of joint consumption of leisure preferences in participation decisions. Controlling for whether the husband has retired is likely to generate simultaneity bias, but his age (relative to his wife's) is a good exogenous proxy for his retirement status. Care responsibilities are represented by variables that measure the numbers of children of different ages as well as the number of parents or parents-in-law reported to need care. Note that a report of needing care does not necessitate that the observed woman functions as the caregiver, as that could introduce simultaneity bias. While adult children (over age 18) are unlikely themselves to require parental time, they proxy for the presence of grandchildren, which otherwise are unobserved in these data (unless they live in the same household as the observed woman and using only those cases would create a selection problem).

The women's own age is measured through a set of binary variables that represent two-year intervals, allowing for a very flexible fit of participation to age. Various trends may be at play here. From a human capital perspective, increased age may represent greater investments in human capital accumulated through on the job training, and therefore lead to higher wage offers. On the other hand, as described in the previous section, some sources suggest that economic reforms have created a work environment that is increasingly discouraging towards older women, thereby counteracting the potential experience effect. Finally, pension eligibility, tied to age, is likely to influence participation decisions of women in covered positions.

The model also includes controls for geographic region (whether city, suburb or town), fixed province effects and fixed year effects. Standard errors are clustered at the level of the individual woman. Marginal effects, calculated from the probit coefficients, and p-values (in parentheses) are presented in Table 1. Turning to the results, we find that a 1,000-*yuan* increase to non-labor income reduces the probability of participation by one percentage point. Participation falls uniformly with age, and the impact grows larger and more significant at each age increment. Women in the age 48–49 category are nearly 30 percentage points less likely to participate, relative to the omitted group of women aged 40–41 years. The oldest category of women, age 50–51 years, experience a 42-percentage point relative decline. Consistent with human capital theory, the marginal effects of schooling are positive with magnitudes rising for each incremental level of achievement.

Considering household demographic factors, a wider gap in age between the husband and wife reduces women's labor force participation, but the effect is very small and statistically significant at only 15 percent. However, the variables representing numbers of children all have negative coefficients, and several are found to be statistically significant. Namely, children below age 7, between ages 7 and 12, and between ages 19 to 25 all display significant negative impacts. As described above, negative effects observed on coefficients

for kids older than age 18 are interpreted to represent grandchild care needs. Within the CHNS data set, among urban women reporting births from 1991 forward, 35 percent of the births of a first child occurred to women age 24 or below, 48 percent occur by age 25, and 60 percent occur by age 26. For men, thirty percent of the births of first children occur by age 25, and half occur by age 27. Therefore, it seems reasonable to infer that observed declines in the labor force participation of the mothers of children in their twenties is associated with the care obligations generated by the arrival of grandchildren.

Though the marginal effects of parent and parent-in-law need for care have a negative sign, their impacts are not significantly different from zero at conventional levels of significance. In contrast to the case of grandchildren, the care needs of the elderly may be relatively more short-lived, ending either with successful resolution (if the elder remains living) or with the elder's death, upon which point there is no longer a need for care. Therefore, while elder care needs may trigger labor force withdrawals, it is possible that the correlation is weak simply because the need for care terminates even as withdrawals persist. This situation could arise if a woman who withdraws from the labor force in response to elder care needs finds it difficult to re-enter when her care obligations have concluded.

In order to shed further light on factors responsible for the timing of women's withdrawals from the labor force, my next step is to estimate the discrete time hazard model over a set of covariates very similar to that described above. Added to the model are indicators of the women's employment sector and occupation (possible because all women are observed to be employed until they withdraw), as well as binary indicators of observed spell length that together form the piece-wise baseline hazard function. Interactions of these interval dummy variables with covariates are included in order to allow for the possibility of non-proportional hazards. The logit specification affords additional flexibility in this regard (Jenkins, 1995). Due to my failure to reject the null hypothesis that random effects are jointly zero, they are omitted from the estimation, but the standard errors are clustered at the level of the individual woman.

Six columns of results are presented in Table 2. The first two columns report the results from a model estimating the hazard of any kind of labor force participation withdrawal, the third and fourth columns examine non-retirement withdrawals, and the fifth and sixth columns examine retirements. The choice between retirement and non-retirement withdrawal is modeled as a competing risk model using a multinomial logit specification. As described in the previous section, withdrawals can be distinguished into two types—retirement (self-reported), and otherwise. The logit coefficients are exponentiated in order that they can be interpreted as log-odds ratios (so a value of one represents a null result). As before, p-values are in parentheses.

Demographic factors along with employment characteristics are shown to have significant influence on the estimated hazards, although different combinations of factors are associated with the different withdrawal types. For instance, women age 48 and above display a withdrawal hazard that is 6 times greater than that of the omitted age 40 to 43 group.² The withdrawal hazard for women in their mid-forties, on the other hand, is only slightly elevated relative to the younger group. Age is not a significant factor in non-retirement

withdrawals, though we see that it is very important for retirements. The retirement hazard for women age 44 to 47 is nearly 3 times greater than that of women age 40 to 43, and it is nearly 20 times greater for women age 48 to 51. These magnitudes simply reflect the important contribution of age towards the likelihood of withdrawal. Rates of withdrawals, and especially retirements, are very low among women in their early forties. But as these estimates show, the likelihoods increase very sharply with age, particularly in the case of retirements. Non-retirements, however, reflect no correlation with the woman's own age.

The age of husbands (relative to their wives) is shown to have only a mild impact, and solely on non-retirement withdrawals; each one year of age difference is equivalent to a 7 percent change in the hazard rate. Non-retirement withdrawal most likely offers women the most discretion with respect to choosing to withdraw in tandem with a spouse. In addition to the joint consumption of leisure factor, care needs may also play a role.

Women's schooling levels and employment characteristics also exert some interesting influences on patterns of labor force withdrawals. Schooling attainments generally correlate with reductions in the hazard rates of any kind of withdrawal, though the effects are significant only for technical or post-secondary schooling. Women in this category experience significantly lower hazard rates across the board, but the magnitude of the effect is largest for non-retirement withdrawals, where the hazard decreases by more than 70 percent. Similarly, having a white-collar occupation (professional, technical, or administrative) reduces all withdrawal hazards but in this case the magnitude of the effect is larger for retirements, where the hazard declines by nearly 80 percent, and the effect is not significantly different from zero for non-retirement withdrawals. This result is consistent with the later official retirement age (55 versus 50) faced by white-collar women.

The employment sector variable is constructed to distinguish between women employed in any public (government) or publicly owned unit (state- and collectively-owned enterprises) versus women who are self-employed, or employed in privately- or foreign-owned enterprises. Key among the differences between these two groups is that women in the government/state/collective sector are far more likely to be covered by the pension system. Predictably, retirement and non-retirement hazards are experienced very differently across these sectors, even though the overall withdrawal hazards are not significantly different between the two groups. This can be explained by the fact that sector effects offset one another—for the government/state/collective group, non-retirement hazards are lower (by at least 47 percent) while retirement hazards are higher (by more than 200 percent). So while overall withdrawal hazards are not significantly higher or lower, in fact those employed by public or publicly-owned work units are much more likely to take retirement rather than non-retirement withdrawal.

Variables that represent the potential dependent care and support needs the women may face also are shown to have important influences. Two specifications of numbers of children are offered, with the first being more disaggregated than the second. The evidence suggests that

²I have collapsed the age variables into three categories each representing 4 years in this estimation exercise. Testing revealed no significant differences in coefficients on the pairs of 2-year variables represented within each 4-year age span.

adult children exert the largest and most significant impact on women's withdrawal patterns. Under the first more disaggregated specification, a child aged 19 to 25 increases the hazard of any withdrawal by more than 40 percent, and this same child increases the non-retirement withdrawal hazard by 75 percent. Tests of significance reveal that the coefficients estimated on the more narrowly defined age categories do not differ significantly from one another within the 18 and below set as well as within the 19 and above set. The results derived from the specification where kids are aggregated by age into only two categories reinforces the idea that adult children serve to increase the hazard of any withdrawal (by 39 percent) as well as the non-retirement withdrawal hazard (by nearly 72 percent). Mirroring the earlier discussion surrounding the probit model results, the finding that children of adult age have the greatest impact on women's withdrawals, as these women move through their middle-age years, underscores the important contribution of grandchild care obligations to women's employment trajectories.

Meanwhile, the elder care needs coefficients presented here differ from the probit labor force participation results in two ways. First, many of the coefficients are statistically significant, and second, the sign of the impact varies depending on whether it is the women's own parents or her parents-in-law who need care. Each parent needing care decreases the hazard of withdrawal by more than 40 percent while each parent-in-law needing care increases the hazard by 60 percent.³ Perhaps a woman's working horizon is extended by obligations to send financial support to her own parents, while traditional obligations to in-laws leads to a greater likelihood of displacement from the labor force. The effect of parents-in-law care needs is particularly important for non-retirement withdrawal hazards, where each additional elder care charge raises the hazard by 80 percent. As with the case of children, the effects of elder care needs on retirement hazards are insignificant.

Interestingly, the income control is insignificant in all models. While non-labor income was found to display a modest negative correlation with labor force participation status, it does not appear to influence the timing of the withdrawal. It is possible that other factors dominate the timing decision. However, it also may be the case that my non-labor income measure (including asset and transfer income) fails to sufficiently capture the variations in financial resources that are most important for withdrawal timing, as the variable takes a zero value for more than two-thirds of the sample. Giles, *et al.*, (2011), using a measure of housing wealth, find a negative correlation with individual's probability of employment for adults over age 45, though it is statistically significant for only the male portion of the sample. Unfortunately, the CHNS data do not offer information on either housing values or household financial assets.

To lend further insight into the pattern of labor force withdrawals over time, I use the coefficient estimates derived above to calculate the predicted hazards of withdrawal for a hypothetical woman, and examine how the hazard rate changes as the woman ages. One case positions a 41 year-old woman in the year 1991, and predicts her withdrawal hazards for the

³Liu, Dong, and Zheng (2010) report a somewhat similar finding where a woman's care provision for parents-in-law reduces both the probability of being employed as well as hours in paid employment while the provision of care for her own parents has no employment impact.

current year and over the 3 subsequent waves (1993, 1997, and 2000), allowing her age and her other demographic data to update accordingly.⁴ The second case proceeds similarly, beginning with the same 41 year-old woman, but places her in the year 1997, and follows her through the 2000, 2004, and 2006 waves. The third and final case places a 41 year-old woman in 2004 and follows her through the 2006 and 2009 waves.⁵ Hazards are constructed for three different events: any withdrawal, a non-retirement withdrawal, and retirement. The predicted hazards are plotted by age, with the different survey wave year cohorts shown in the same diagram, in order to illustrate how withdrawal patterns, by age, have evolved over the twenty-year span of data. Figures 5, 6, and 7 display these results.

For a woman entering her forties in the early 1990s, the hazard rate of withdrawal (Figure 5) is relatively low at the outset and increases incrementally as she ages. In contrast, for a woman entering her forties in 1997, the initial withdrawal hazard is much higher, being double, at 40 percent, that of the 1991 cohort counterpart. By the time she reaches age 44 (in the year 2000), her withdrawal hazard has risen to higher than 60 percent. It seems likely that economic restructuring in the latter 1990s and early 2000s contributes to high predicted rates of withdrawal among women in this cohort who might otherwise be considered to be within their prime working years. The interpretation is confirmed by the pattern displayed for the 2004 cohort where withdrawal hazards seem to be returning to, or below, the earlier levels.

Comparing non-retirement to retirement hazards, there are two noteworthy observations. First, the invariance of non-retirement hazards to age (in Figure 6) contrasts sharply with the sharp increases, by age, observed for retirement hazards across all cohorts (in Figure 7). Second, while retirement hazards appear to be declining with each successive cohort (again in Figure 7), non-retirement hazards appear to be rising, especially for women in their early to mid-forties (Figure 6). So while labor force withdrawal patterns, overall, may have remained reasonably steady (setting aside the restructuring episode), the composition of withdrawal types has changed. The data suggest that hazards of non-retirement withdrawal have increased over time while retirement hazards fallen.

V. Pension receipt and income

Among the 238 women observed to withdraw from the labor force, those who take non-retirement withdrawals uniformly report no receipt of pension income, while most (82 percent) of those who retire indicate some receipt of pension income in the subsequent survey wave. As such, only a little more than half of the withdrawn women (53 percent) report earning a pension. Table 3 provides further description of women's pension receipt, by occupation and sector of employment.

⁴For this prediction exercise, I replace the set of age category binary variables with age and age squared in order to pinpoint the specific hazard rate for the given age. Completed schooling is set at the lower secondary level. Neither the public employment sector nor the white-collar occupation dummy variables are turned on. At age 41, I assume the woman has one child age 18 or below, and this child ages to over 18 by the time she is 43. The elder care variables are always turned off. The woman is assumed to live in Shandong province. Continuous variables are set at sample means.

⁵I cannot predict withdrawal outcomes for the 2011 wave because the underlying regression estimation takes the withdrawal outcome as observed in the subsequent wave, and 2011 is the final wave in the data sample.

Holding a white-collar position adds some advantage to the likelihood of pension receipt among the full withdrawn sample; 63 percent of the white-collar group report a pension relative to 50 percent of those in other occupations. With respect to employment sector, the self-employed are by far the least likely to receive pension income, with receipt at only 11 percent of the full subset of the withdrawn self-employed. This is to be expected, given reports that the self-employed have been the slowest to enroll in pension programs (Giles, *et al.*, 2013). The variation within remaining sectors is relatively modest, though the government and collective sectors offer the best chances of pension income (65 percent in each case) while the private and foreign sectors offers the worst (47 percent).

In order to understand how pension income may contribute towards women's financial positions, I examine the level of pension income received in the survey wave subsequent to withdrawal (deflated to constant 1988 *yuan*) as well as its percentage of the net individual income reported in the preceding survey wave when the woman was still working.⁶ Reported pension income levels are highest among the self-employed, but ironically this group has the lowest ratio of income replacement, at only 26 percent, reflecting the extraordinarily high incomes earned, on average, among this group. At the other end of the spectrum, pensions are lowest for women in the state enterprise sector, although here we see an income replacement ratio of 117 percent, reflecting the low earnings in this sector. Women retiring from white-collar positions benefit from somewhat higher pension levels that also do a better job of replacing their former incomes. In general, the replacement ratio is high, at more than 100 percent for the full group of pension-earners.⁷ However, concerns are raised by the considerable portion of women who are not receiving pensions, which constitutes nearly half (47 percent) of those who have withdrawn, giving no indication of subsequent labor force re-entry.

Pensions form the primary source of income for the elder population in urban China (He and Sato, 2013; Park, *et al.*, 2012). Additional sources support for the elderly are earnings (if they remain employed), their savings, and private transfers. Public transfer payments, such as the minimum living allowance guarantee, are less important (Park, *et al.*, 2012). This points to a particular vulnerability of women who fall outside the pension system and are called upon to withdraw from the labor force during their prime earning years, as re-entry may be difficult, thereby compromising their future employment and earnings opportunities. While non-participation in the pension system may allow women to be more responsive to caregiving needs, it no doubt also generates an even greater dependency on family, in the long run, to offset the lost earnings and savings that continuous employment could provide. Not surprisingly, Chen and Turner (2015) report that the portion of older women reliant on family support (53 percent) is nearly double that of older men (28 percent).

⁶In order to avoid undue influence from a few extreme values in reported income (particularly among the self-employed), I use the median levels of reported pension and reported income.

⁷It is worth noting that the replacement ratio may be biased upwards somewhat because the pension amount is taken from the wave year subsequent to the income measure. Taking real wage growth into account, one might expect that the woman's income would have increased had she remained in the workforce, yielding a lower replacement ratio of pension to foregone income. Thanks to an anonymous referee for raising this point.

VI. Conclusions

My results give way to several policy recommendations. First, increasing the retirement and pension eligibility age can reliably extend the working years of women who participate in the pension system. Support is offered by evidence showing that retirement hazards are extremely sensitive to age, and that women in white collar occupations (with later official retirement ages) experience both lower overall withdrawal hazards and lower hazards of retirement. However, a later retirement age will not motivate later withdrawals for the many women who are not covered by the pension system. I find that non-retirement withdrawals are relatively invariant to age and constitute a rising portion of the labor force withdrawals observed among women in my sample. Therefore, in order to address China's concern about a declining labor supply, the pension system needs to increase its breadth of coverage in order that the age lever can incentivize longer working years among a larger portion of the work force.

Making the system more inclusive offers additional advantages. Almost half of the women observed to have withdrawn from the labor market in my sample report no pension receipt. This deficiency raises concern about women's capacities to provide for themselves in old age. I find that women with private sector or self-employment are the most at risk, and these results are consistent with reports that pension system participation is lowest in these sectors. Of course, broadening participation also increases the system viability by spreading risk across a larger pool of individuals. Many have advocated for a more unified centrally-administered system as a first step towards establishing a comprehensive national pension plan (see for example He and Sato, 2013; Pozen, 2013; or Cai and Cheng, 2014), and my results add support for this policy action.

Better institutional support for dependent care needs (such as government sponsored child and elder care) also is warranted because I find that family considerations play an important role in the timing of non-retirement withdrawals. Older husbands confer higher non-retirement withdrawal hazards on their wives. Adult children (proxying for the existence of grandchildren) and parent-in-law need for care also raise non-retirement withdrawal hazards. These caregiving demands may escalate in coming years. A woman in the one-child policy cohort may be the only potential caregiver for both her parents and her parents-in-law. As this cohort of women age into their forties and fifties, we can expect an even higher incidence of labor force withdrawals driven by elder care needs. Ironically, the recent reversal of the one-child policy, allowing couples to have two children, may compound the pressure on middle-aged women by increasing their grandchild care obligations.

Finally, eliminating interruptions and lengthening women's career horizons stimulates a virtuous cycle by generating incentives for human capital investment, which subsequently reinforce women's labor force attachments throughout their working lives. Indeed, my results show that technical and post-secondary schooling reduce women's hazards of labor force withdrawal. Therefore, we can expect larger schooling investments to be accompanied by longer employment trajectories.

Supplementary Material

Refer to Web version on PubMed Central for supplementary material.

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Figure 1.
Labor force participation rates for urban married women aged 25–64, by age and year (4,594 observations)
Source: Author's calculations using CHNS data

Figure 2.
Labor force withdrawal rates of urban married women aged 40–51, by age (1,076 observations)
Source: Author's calculations using CHNS data

Figure 3.

Non-retirement withdrawal rates of urban married women aged 40–51, by age (1,076 observations)

Source: Author's calculations using CHNS data

Figure 4.
Retirement rates of urban married women aged 40–51, by age (1,076 observations)
Source: Author's calculations using CHNS data

Figure 5.

Predicted withdrawal hazards, by age and cohort

Source: Author's calculations using CHNS data and IMF *International Financial Statistics*

Figure 6.

Predicted non-retirement withdrawal hazards, by age and cohort

Source: Author's calculations using CHNS data and IMF *International Financial Statistics*

Figure 7.

Predicted retirement hazards, by age and cohort

Source: Author's calculations using CHNS data and IMF *International Financial Statistics*

Table 1

Probit estimation of labor force participation

Labor force participation	
Dependent variable mean	0.672
Independent variables	
Non-labor income (thousand 1988 <i>yuan</i>)	−0.011 ** (0.01)
Age 42–43	−0.035 (0.22)
Age 44–45	−0.057 * (0.08)
Age 46–47	−0.148 *** (0.00)
Age 48–49	−0.298 *** (0.00)
Age 50–51	−0.423 *** (0.00)
Primary school attainment	0.102 *** (0.01)
Lower secondary school attainment	0.200 *** (0.00)
Upper secondary school attainment	0.255 *** (0.00)
Technical or postsecondary	0.384 *** (0.00)
Husband's age difference (years)	−0.005 (0.15)
Number of kids age 6 or below	−0.171 ** (0.02)
Number of kids age 7 to 12	−0.063 * (0.08)
Number of kids age 13 to 18	−0.026 (0.26)
Number of kids age 19 to 25	−0.033 * (0.09)
Number of kids age 26 and above	−0.045 (0.16)
Number of parents needing care	−0.032 (0.13)
Number of parents-in-law needing care	−0.011

Labor force participation	
	(0.65)
Resident of suburb	0.004
	(0.89)
Resident of town	0.043
	(0.11)
Fixed province effects	yes
Fixed year effects	yes
Observations	3,136
Number of observations reporting participation	2,108

Robust pval in parentheses

p<0.01

**
p<0.05

*
p<0.1

Notes: The table displays marginal effects calculated from the probit coefficients at the mean values of the independent variables

Source: Author's calculations using CHNS data and IMF *International Financial Statistics*

Table 2

Discrete time hazard estimation of labor force withdrawals

	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable means	0.221	0.221	0.078	0.078	0.143	0.143
Independent variables						
Non-labor income (thousand 1988 yuan)	1.037 (0.51)	1.036 (0.51)	1.070 (0.23)	1.069 (0.24)	0.998 (0.98)	0.999 (0.98)
Age 44–47	1.480* (0.10)	1.480* (0.10)	0.958 (0.90)	0.945 (0.87)	2.777*** (0.00)	2.797*** (0.00)
Age 48–51	6.200*** (0.00)	6.095*** (0.00)	0.978 (0.97)	0.961 (0.94)	19.27*** (0.00)	18.82*** (0.00)
Husband age difference (year)	1.029 (0.30)	1.028 (0.31)	1.072* (0.06)	1.070* (0.07)	0.986 (0.69)	0.985 (0.66)
Primary school attainment	0.550 (0.19)	0.547 (0.18)	0.542 (0.31)	0.527 (0.29)	0.606 (0.34)	0.606 (0.35)
Lower secondary school attainment	0.884 (0.77)	0.872 (0.75)	0.640 (0.38)	0.624 (0.36)	1.125 (0.82)	1.107 (0.85)
Upper secondary school attainment	0.567 (0.20)	0.561 (0.19)	0.460 (0.16)	0.443 (0.14)	0.709 (0.53)	0.706 (0.52)
Technical or post-secondary	0.406* (0.08)	0.405* (0.08)	0.270* (0.08)	0.260* (0.07)	0.544 (0.33)	0.546 (0.33)
Professional/technical/administrative/manager	0.381*** (0.00)	0.384*** (0.00)	0.597 (0.19)	0.601 (0.19)	0.237*** (0.00)	0.241*** (0.00)
Government/state/collective sector	0.664 (0.16)	0.663 (0.16)	0.424** (0.02)	0.427** (0.02)	3.147** (0.03)	3.111** (0.03)
Number of kids age 6 or below	0.931 (0.92)		1.229 (0.82)		0.657 (0.70)	
Number of kids age 7 to 12	1.030 (0.93)		1.836 (0.16)		0.654 (0.43)	

	(1)	(2)	(3)	(4)	(5)	(6)
	Any withdrawal	Any withdrawal	Non-retirement withdrawal	Non-retirement withdrawal	Retirement withdrawal	Retirement withdrawal
Number of kids age 13 to 18	0.959 (0.84)		1.389 (0.27)		0.686 (0.15)	
Number of kids age 19 to 25	1.417** (0.04)		1.753** (0.03)		1.168 (0.46)	
Number of kids age 26 and above	1.224 (0.62)		1.555 (0.53)		0.936 (0.88)	
Number of kids age 18 or below		0.963 (0.85)		1.420 (0.21)		0.673 (0.12)
Number of kids age 19 and above		1.393** (0.05)		1.719** (0.04)		1.133 (0.54)
Number of parents needing care	0.663* (0.08)	0.666* (0.08)	0.680 (0.27)	0.677 (0.26)	0.672 (0.19)	0.680 (0.20)
Number of parents-in-law needing care	1.605** (0.03)	1.596** (0.03)	1.830** (0.04)	1.790** (0.04)	1.437 (0.17)	1.431 (0.17)
Resident of suburb	0.824 (0.42)	0.823 (0.42)	0.784 (0.56)	0.778 (0.55)	0.995 (0.99)	0.993 (0.98)
Resident of town	0.929 (0.73)	0.935 (0.76)	1.421 (0.30)	1.429 (0.29)	0.674 (0.18)	0.678 (0.19)
Constant	0.148*** (0.01)	0.153*** (0.01)	0.039*** (0.00)	0.042*** (0.00)	0.029*** (0.00)	0.031*** (0.00)
Fixed province effects	yes	yes	yes	yes	yes	yes
Fixed year effects	yes	yes	yes	yes	yes	yes
Piecewise baseline hazard and interactions	yes	yes	yes	yes	yes	yes
Observations	1,076	1,076	1,076	1,076	1,076	1,076
Number of observations reporting a withdrawal	238	238	84	84	154	154

Robust pval in parentheses

p<0.01**
p<0.05

Notes: The table displays log-odds ratios derived by exponentiating the logit coefficients
Source: Author's calculations using CHNS data and IMF *International Financial Statistics*

$t < 0.1$
*

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Table 3

Pension receipt by occupation and ownership

	Percent receiving a pension		Relative pension amount		Number of observations
	Among retirees	Among all withdrawn	Level in constant 1988 <i>yuan</i>	Replacement ratio	
Professional, technical, administrative or management	0.80	0.63	3133.54	1.05	52
Office, production or service staff	0.82	0.50	2165.89	1.01	186
All occupations	0.82	0.53	2403.66	1.09	238
Self-employed	0.67	0.11	3175.29	0.26	38
Government or institute	0.82	0.65	2307.72	0.99	124
State-owned enterprise	0.76	0.55	1882.40	1.17	40
Collectively-owned enterprise	0.92	0.65	3032.24	0.96	17
Private or foreign-owend enterprise	0.90	0.47	2741.60	0.89	19
All ownership types	0.82	0.53	2403.66	1.09	238

Source: Author's calculations using CHNS data and IMF *International Financial Statistics*