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A Unit Record Analysis of Older Male Labour Force Participation

Martin O'Brien

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This paper presents empirical analyses of econometric models of older males' labour force participation based upon the orthodox theory of labour supply. The aim is to assess the effectiveness of using of micro data to account for older male labour force participation rate patterns over recent decades. As such, the influence of financial variables from the budget constraint and observable characteristics from the utility function are incorporated into reduced form models estimating the older male's expected retirement age and the probability of an older male's labour force participation using unit record data. The research findings presented emphasise the role of human capital and potential market wages for an older individual's labour force participation. In contrast, the coefficients attached to observable characteristics variables, representing tastes and preferences for work and leisure, were rarely statistically significant or of a consistent sign across age groups or time periods, meaning that that unobservable tastes may dominate the labour force participation decision. A major finding was that higher wage earners were more likely to anticipate an earlier exit than lower earners. This shows that expectations do not necessarily match actual labour force participation behaviour, and that lower wage earners may be more vulnerable to unanticipated labour force exit. Finally, micro-based findings were not adequately translated to the macro level, being unable to directly offer inferences for aggregate changes in older male labour force participation rates over time. This problem of aggregation results in a fallacy of composition in reasoning, meaning that the micro models presented have only limited use for explaining older male labour force participation rate patterns over recent decades.

* School of Economics and Information Systems, University of Wollongong.
Email: martinob@uow.edu.au

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1. Introduction

A number of recent Commonwealth government policy reforms have sought to increase the labour force participation of older workers aged 55 to 64 years (for example DFaCS 2000, 2001, 2002a, 2002b, and House of Representatives 2000). As such, recent attempts have been made to restrict access to various social security pensions that have traditionally accommodated older males not in the labour force. Other policies put forth are designed to encourage training and non-standard employment. Policy can be largely characterised as supply side, affecting the individual's budget constraint variables and (labour) marketable skills. For example, no employment subsidies or direct job creation policies have been foreshadowed.

However, empirical research into factors affecting older worker labour force participation remains rather scarce in Australia, most being conducted in the 1980's and is now quite dated.[1] In comparison, empirical research in the US has continued and developed from the late 1960's to now, increasing structural and dynamic complexity.[2] These estimated models generally use unit record data, chosen from a multitude of available longitudinal survey datasets. Empirical models are generally a depiction of the orthodox model of labour supply. That is, emphasising the individual's labour supply reaction to the relative prices of work and leisure (wage and non-wage income variables). Results have generally been used for aggregate labour force participation rate forecasts and various policy simulations.

In comparison to the US literature the majority of the (dated) Australian empirical research analysed aggregated labour force participation rate data. It is therefore of particular concern that policy reforms have been directed at the individual when the microeconomic research is poorly understood in Australia. With a number of Confidentialised Unit Record Files (CURFs) becoming available in Australia over recent years it would appear an opportune time to assess the adequacy of unit record data in explaining and predicting older male labour force participation in Australia. Thus, this paper presents and assesses a number of unit record econometric models estimating the probability of an older male's labour force

participation as a function of their financial and demographic characteristics. An assessment is then made as to the worthiness of these micro model results accounting for prevalent older male aggregate labour force participation rates over time, as has been used in the US literature. The findings are intended to enhance the understanding of older male labour force participation behaviour and have a number of implications for supply side policy makers.

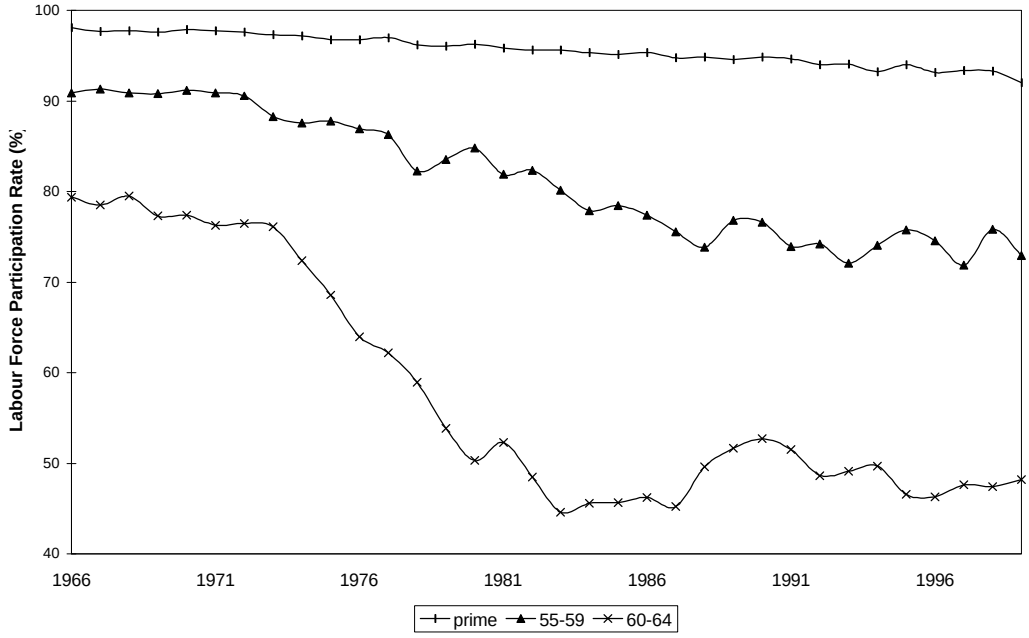
The structure of the paper is as follows. Section 2 presents the trends in older male labour force participation rates over recent decades in Australia. The theoretical framework is also explored, differentiating between orthodox, structural labour market, and aggregate labour market explanations. An exploration of overseas and Australian empirical literature explaining older male labour force participation is reviewed in Section 3. The modelling of unit record data and use of longitudinal data in the US research is particularly highlighted. Section 4 outlines the econometric model specification and estimation method. Sections 5 to 7 contain estimation results for the binomial labour force participation model, multinomial labour force participation model, and retirement expectation models, respectively. Section 8 assesses the adequacy of the micro theory and model results in accounting for prevailing labour force participation patterns over time. Conclusions are presented in Section 9.

2. Trends in older male labour force participation and theoretical framework

A marked trend in Australian and international labour markets over recent decades has been the decline in labour force participation rates of older males (for example, O'Brien 1999, and Stille 1999). In Australia at least, this trend is in contrast to generally increasing labour force participation rates for older females (O'Brien 2000). It is shown in Figure 1. that while there has been a decline in labour force participation rates over time for all male age groups presented, this decline appears to increase with age group. The decrease in labour force participation rates from 1966 to 1999 is equal to six, eighteen and thirty one percentage points for males prime-aged (aged 25-44 years), 55-59, and 60-64 years, respectively. The rapid decline in labour force participation rates of older males began in the early 1970s and extended into the 1980's. A minor rebound is evident in the late 1980's.

However, older male labour force participation rates have remained at their relatively lower (1980s) rates throughout the 1990's.

Figure 1 Australian male labour force participation rates (%) – 1966 to 1999



Source: ABS (1986, 1999a)

The theoretical framework to account for these trends in older male labour force participation rates can be divided into three broad explanatory groups. In turn, these three explanatory groups are differentiated by the unit of analysis required to address suggested testable propositions. First, the orthodox theory of labour force participation emphasises the role of relative prices and preferences for work and leisure faced by the utility maximising individual and should be analysed at the unit record level. Second, structural labour market explanations utilising neoclassical labour demand or early generations of segmented labour market theory emphasise the role of the employment relationship and the institutional framework for the older worker and are best analysed at the firm or industry/occupation level. Third, aggregate labour market explanations such as labour force discouragement and

the Reserve Army of Labour (RAL) hypothesis are best analysed at the macroeconomic level.

The familiar (static) orthodox model of labour supply consists of an allocation of time between paid work and leisure for an individual, and is derived from the general model of consumer demand (Hicks 1946). This individual seeks to maximise utility subject to the budget constraint. That is,

$$\text{Max } U = (x, h; A, e) \tag{1}$$

subject to

$$px = wh + y \tag{2}$$

The utility function for the individual (equation 1) is defined over consumption of commodities, x, and hours of work h, and varies across individuals according to personal characteristics A (observed) and ε (unobserved) which affect tastes and preferences for work and leisure. This individual is constrained in the pursuit of utility maximisation and commodity consumption (the product of prices and quantity of commodities, *px*) by the resources they possess. That is, the budget constraint (equation 2). In turn, the budget constraint is derived from income received from paid employment in the labour market (product of wage rate and hours worked, *wh*), and non-wage income (*y*).

The primary interest in this study is the labour force status, rather than the specific number of hours supplied to the labour market. Therefore, one would not necessarily be interested in substitution and incomes effects from a change in market wage (the Slutsky equation) unless it led to a change in labour force status. The key for labour force participation is to distinguish between *interior* and *corner* solutions to the constrained maximisation problem. For an interior solution, first order conditions for the constrained maximisation problem require that *x* and *h* are chosen so that the (negative) value of the marginal rate of substitution of working hours for commodities (*m*) equals the real wage (*w/p*). That is:

$$\frac{w}{p} = \frac{dU/dh}{dU/dx} = -m(x, h; A, e) \qquad (3)$$

Stated differently, at this level of hours worked, the rate at which the individual is prepared to substitute leisure for work in order to maintain the specified utility level (the marginal rate of substitution) is equal to the market valuation of work (the real wage). For a corner solution, the marginal rate of substitution exceeds the real wage at all possible hours of work.

This static model of labour force participation can be extended in a number of ways. First, to incorporate non-linear budget constraints and restrictions on hours of work imposed by employers.[3] Second, the model can also be extended with the endogenisation of the market wage profile with the incorporation of human capital theory, emphasising the skills possessed by the individual, and the role played by investments in education, and on-the-job training (Mincer 1974). Third, the lifecycle model of labour force participation provides an extension of the static one period model into a dynamic multi-period analysis of an individual's labour force participation (for example Ghez and Becker 1975).

In summary, the static orthodox model of labour force participation, which forms the foundation of subsequent extensions to the orthodox labour force participation model, is based on two fundamental principles. First, the individual's tastes and *relative preferences* for work and leisure, represented by observable characteristics A, and unobservable characteristics e, in their utility function. Second, individuals react to financial incentives affecting the *relative prices* of work and leisure, represented by wage (w) and non-wage income (y) in the budget constraint. The individual, being the decision making unit, *chooses* his/her ultimate labour force status and hours of labour supply according to these financial constraints faced. The external environment such as labour market opportunities is only indirectly incorporated into the budget constraint faced by the individual, and thus exogenous and not the focus of the model. This has the implication that any resulting labour

force participation decision is the result of a voluntary action of the individual, the outcome reflecting the individual maximising their utility according to the options faced at all times.

The orthodox model of labour force participation can potentially be used to explain changes in older male labour force participation rates over time through changes in the relative work/leisure prices faced (w and y) or the observable and unobservable characteristics of the older male population (A and e), respectively affecting aspects of the budget constraint and utility function. However, there are two fundamental problems aggregating inferences based upon this microeconomic model to explain the older male labour force participation rates presented. First, aggregate labour force participation inferences based upon this theory should relate to older male employment to population rates, rather than labour force participation rates, noting the theoretical distinction between positive versus zero hours. Second, and more fundamental, is that the orthodox model of labour force participation is based upon the microeconomic unit of analysis and does not directly provide aggregate labour force participation inferences, without imposing restrictive assumptions.[4]

3. Empirical literature and data availability

Empirical research on older male labour force participation from the U.S., and specifically models inspired by orthodox labour force theory, dominate the literature.[5] Researchers such as Boskin (1977) and Parsons (1980) were some of the earliest researchers to estimate the probability of older male workers' labour force participation. Generally using a binomial logit estimation on cross sectional unit record data, the roles of non-wage (social security) eligibility and value, bad health, and job attributes were established in early reduced form static model specifications. Later models have incorporated private pension value and eligilibility rules (Stock and Wise 1990), as well as early retirement pension offers and windows of opportunity (Hogarth 1988, Lumsdaine et al 1990). In general, econometric models of older males' labour supply became increasingly *dynamic* and *structural* over time, largely reflecting aspects of the lifecycle model and exploiting the longitudinal nature of the datasets available. In contrast to the static model depiction, *dynamic* models incorporate

lifecycle models concepts through incorporation of the present value of variables (Burkhauser 1980), and can allow some of the model's parameters to be updated as time passes and new information is incorporated (Rust 1989). In contrast to reduced form specification, [6] *structural* models are specified in terms of the theoretical conception, and as such incorporate the estimation of utility functions and detailed aspects of budget constraints. [7] As such they allow the estimation of theoretical parameters. This necessarily entails an increase in complexity and generally assumes a higher level of rationality on behalf of the individual. However, even with the increase in technical complexity the same core findings are generally as before. That is, findings still emphasise participation an individual's labour force participation timing as a reaction to financial variables. Recent arguments are largely over the relative size of social security versus private pension financial incentives for retirement, and which particular technique to use. A major implication is that prevalent labour force participation trends can be reversed by the reversal of early retirement pension incentives.[8] The US empirical model development has obviously been aided by the increasing abundance of detailed longitudinal databases allowing structural dynamic model specification. These include the National Longitudinal Survey, Panel Study of Income Dynamics, Retirement History Study, and Health and Retirement Study.

Most Australian older male labour force participation empirical research was published in the early 1980's, with researchers primarily focused on the then recent decline in labour force participation of older men experienced in the 1970's. Similar to U.S. findings, the consensus that emerged emphasised the operation of the orthodox theory of labour force participation. Merrilees' (1986: 220) summarised the literature of competing theories of the economic determinants of older male labour force participation from this era. He claimed:

> In summary, a **consensus** is at last emerging which emphasises the life cycle private wealth effects as the key to the long term trend to early retirement, though not necessarily the short run (1973-76) acceleration in such. (emphasis added)

The lifecycle model under conditions of certainty was used to explain the long-run trend decline in older male labour force participation rates, emphasising the role of increased

wealth and earlier retirement expectations (Hughes 1984, Miller 1983). Increased social security pension payments (Merriless 1982, 1983) and discouraged worker (Stricker and Sheehan 1982) explanations were used to account for short-run fluctuations in older male labour force participation rates. This consensus was therefore largely consistent with orthodox explanations that were favoured in the U.S. older worker labour force participation literature.

A number of methods were used within this consensus era of research such as expected retirement age projections (Miller 1983), linear trend extrapolations of labour force participation rates (Stricker and Sheehan 1982) and econometric modelling (Merrilees 1983, BLMR 1983). In contrast to US research, the majority of Australian research used aggregate data as the unit of analysis. However, an exception, Woodland (1987) addressed the microeconomic determinants of retirement and employment behaviour using unit record data from cross sectional surveys from individuals residing in capital cities in 1981. His findings emphasised the role of pension eligibility, expectations and voluntary behaviour for retirement, while dismissing sectoral and aggregate labour market forces.

A major limitation in the pursuit of unit record analysis of older male labour force participation in Australia has traditionally been the lack of suitable data. Ideally, the empirical analysis of testable propositions from the orthodox labour force participation theory would require the estimation of a structural model using unit record longitudinal data capturing individuals' market wage, reservation wage, and non-wage income value and eligibility from their budget constraint, and variables representing expectations, tastes and preferences for work and leisure over time from their utility function. The first data restriction encountered is that direct information on individuals' expectations and preferences for work and leisure is usually not collected in relevant labour force related surveys, and may only be inferred from the individual's observable characteristics, under the assumption that individuals with similar characteristics share similar preferences (*A* from the utility function in equation 1). Next, unlike the U.S., detailed longitudinal data on individuals' labour force participation patterns is relatively scarce in Australia. Furthermore,

the longitudinal data sources that are available in Australia appear to be unsuitable for the study of older male labour force participation. For example, both the Australian Longitudinal Survey[9] and the Survey of Employment and Unemployment Patterns (SEUP)[10] are targeted more toward youth employment transitions. Finally, DFaCS social security recipient longitudinal datasets are not publicly available. Furthermore, as the DFaCS data does not contain information on those not receiving social security payments, any resulting estimates may be subject to sample selection biases. A relatively new longitudinal survey, The Household Income and Labour Dynamics in Australia (HILDA) survey, designed and managed by the Melbourne Institute of Applied Economic Research, has recently become available (Melbourne Institute 2002). This may become an important tool for the analysis of older workers in the coming years. However, limited relevant information is presently available. Anumber of ABS CURFs have become available in recent years. The Survey of Income and Housing Costs (SIHC), Censuses, and Labour Mobility Surveys provide cross sectional snapshops at CURF level at various points over the 1980's and 1990's. The Social Mobility in Australia Project 1973 (SMA) survey from the SSDA was also chosen for review.

All data sources reviewed contain basic labour force status, income and demographic information. The major deficiency from most sources relates to the lack of information on both wage and non-wage income. The Censuses, Labour Mobility Surveys and SMA contain only one income variable. Multiple sources of income are available only from the SIHC. Furthermore, all surveys lack information on potential employment for those not in the labour force (NILF) or unemployed that would allow the calculation of potential market wage rates for these individuals. Limited information would be available for the non-employed in Labour Mobility Surveys if their job ceased in the last 12 months. Past occupation or industry of employment is, however, available from SMA. Furthermore, apart from SIHC, most surveys lack information on other important observable characteristics variables such as mortgage or other loan commitments that would be expected to affect the preference for work and leisure. Similarly, the surveys reviewed generally did not contain information on non-wage eligibility for social security or private pensions. Furthermore,

apart from the SMA, no data source contains retirement age expectation information. SIHC and Censuses would appear to be largely equivalent. Two SIHCs (1985-86 and 1997-98) were chosen primarily because of the richer data for income and financial commitments. The SMA was chosen to offer a comparison for the 1970's.[11]

Therefore, the paucity of longitudinal data on older male labour force participation in Australia compromises the potential for empirical analysis. In common with the earlier U.S. empirical work (for example Boskin 1977, Parsons 1980), cross sectional analysis will be used, which necessitates that a number of simplifying assumptions be made. The most obvious limitation is that the analysis is restricted to being a static model depiction of labour force participation. This means that analysis of Australian data sources will not be as rich, or as technically ambitious as the studies of U.S. longitudinal data. With this caveat in mind, broad inferences of the suitability of an empirical model based upon orthodox theory to explain older male labour force patterns should still be available to capture the key testable propositions.

4. Labour force participation econometric model specification

The probability of an older male's labour force participation is modelled as a function of budget constraint and utility function variables. The general specification of the labour force participation econometric model is to include the individual's labour force status as the dependent variable. Market wage (w) and non-wage income (y) explanatory variables are specified to represent the relative price of leisure and work from the budget constraint (equation 2). Finally, demographic characteristics and financial commitments are specified to capture observable characteristics (A) and thereby the relative preferences for work and leisure from the utility function (equation 1). That is:

> *Probability of older male's labour force status* = f [*market wage income* (w), *non-wage income* (y), *observable characteristics* (A)]

The model is thus an *ad hoc* reduced form depiction of the theoretical labour force participation decision. However, whereas the orthodox theory is presented in terms of employment ($h>0$) versus non-employment ($h=0$), labour force participation rate trends are presented in terms of the labour force framework of the economically active population of the employed ($h>0$) and unemployed ($h=0$ but desired $h>0$) versus those NILF ($h=0$). This anomaly between theory and the labour force framework was not addressed in previous US or Australian binomial logit based research reviewed. Two types of models are thus estimated, incorporating different dependent variable specifications. First, the binomial probability model estimated in Section 5.4 includes the dependent variable as a dichotomous (0,1) specification of employment versus non-employment. This specification was chosen to be consistent with the labour force participation choice depicted by theory, whereby unemployment and NILF are operationally equivalent non-employment states where the individual's reservation wage exceeds the potential market wage. Estimation results are thus interpreted as predicting the probability of supplying positive hours to the labour market. Second, a multinomial dependent variable specification (0, 1, 2) is used incorporating all three labour force states of employment, unemployment and NILF status allowing the separation of unemployment and NILF labour force statuses on qualitative, if not quantitative, grounds. Of interest is the empirical exploration of any differences between those supplying zero hours to the labour market but actively searching (the unemployed) and those supplying zero hours and not actively searching for employment (NILF).

A key variable explaining the probability of employment is the market wage earnings (w) of the individual. According to the budget constraint, this variable should be distinguished from non-wage income (y) in the labour force participation model specification. However, those unemployed and NILF individuals only receive non-wage income. Wales and Woodland (1980) demonstrate that the ignorance of this issue may lead to estimation biases, and they discuss a number of alternative methods to avoid this bias including proxy variables, sample restrictions, estimation methods and multi-equation specification. In empirical Australian studies, Woodland (1987) did not collect any direct information on market wages income. However, the current occupation for those employed, and the past

occupation for those unemployed or NILF, was available. Woodland then used published ABS occupation wages to proxy for each individual's actual or potential market wages. Alternatively, Borland (1995) imputed labour market earnings for males (of all ages) outside of employment with econometric estimates using human capital proxies (educational attainment and age). The sample restrictions used in U.S. literature to overcome this issue are only available for longitudinal datasets, where the labour force transition of those working in a base period can be subsequently traced.

As past employment occupation or industry is not available in SIHC, a method similar to that used by Borland (1995) is used in this chapter. That is, educational attainment and formal training variables are included in the labour force participation functions to proxy for market wages. This human capital proxy for market wages is a convenient specification as the information is available for older males both in and out of the labour force in the data sources used.[12] The assumption that the education attainment variables may proxy for market wages appears justified with reference to Table 1. There is a clear progression of mean market wages of those employed according to human capital attainment in the SIHC data. Those with a university degree on average possess the highest wages, followed by those with other (certificate and diploma) qualifications, Trade qualifications, then those with no formal qualifications. It is assumed that (potential) market wages for those unemployed or NILF follows the same progression. Therefore, rather than any market wage information directly entering the equations, separate dummy variables (0,1) are included for those older males with university degrees (DEGREE), Trade qualifications (TRADE), and other diploma or certificate qualifications (OTH), with those without formal education or training qualifications forming the control group. The sign of these coefficients is *a priori* ambiguous, depending upon the relative size of substitution or income effects. It is conceded that human capital variables may also capture effects from observable characteristics rather than solely capturing the market wage relation. However, this method was chosen in order to combat the potential sample selection bias problems from the absence of wage information for those unemployed or NILF, which is deemed to be more serious.

Table 1. Mean wage by education or formal training attainment ($ p.a.)

Highest Education /Formal training Attainment	1985-86 Mean Wage ($)	1997-98 Mean Wage ($)
Degree	32676.74	54668.44
Other qualifications	27909.38	47266.54
Trade	21541.07	34595.81
Other	19494.57	27661.11

Source: SIHC 1985-86, 1997-98

Two non-wage income variables are also specified. The first of these, INDIVYQ, is calculated as the difference between the individual's total annual income and market wage income for those employed, and as total income for those unemployed or NILF in both SIHCs. This non-wage information was then ranked and allocated into quintiles. This process was not so straightforward for the SMA, as only one income variable was measured. Therefore, industry wage information from 1973 (ABS 1992) was used as the market wage proxy for those employed, and was subtracted from total income. This non-wage income was then ranked and allocated to quintiles. As with the SIHC, total income in the SMA is used for non-wage income for those unemployed and NILF. The second non-wage variable, UNITYQ, represents the non-wage income available from the rest of the household, including the spouse. This is simply calculated as the difference between total household and individual income, and then allocated to quintiles as above. Whereas the effect of market wages upon labour force participation is *a priori* ambiguous, it is expected that there will be a negative relationship between the level of non-wage income and the probability of labour force participation. Unfortunately, non-wage eligibility for social security or private pensions is not captured in the data sources and cannot be depicted.

Whereas the above explanatory variables capture the *relative price* of work and leisure with *w* and *y* from the budget constraint, the following explanatory variables capture the *relative preferences* for work and leisure from observable characteristics *A* in the utility function. It is implicitly assumed that those individuals with similar observable characteristics share similar preferences for labour force participation (Pencavel 1986), with unobservable preferences left to the random disturbance term (e). The inclusion of marital status (MARITAL) implies that labour force participation may be a joint decision (see Section 2.4.1). This would have an ambiguous effect on labour force participation in theory, however, stylised facts suggested that married males had a greater attachment to the labour force and employment. The presence of dependents (DEP) is also included, however, this is expected to affect financial commitments and is expected to be positively related to labour force attachment. Similarly, paying off a mortgage or other loans is expected to add to the individual's financial commitments and the need for an income from employment (LOANS).[13] In contrast, home ownership (HOWN), that is having paid off the mortgage, is expected to relax the financial commitments and possibly allow labour force exit.

A summary of explanatory variable specification for older male labour force participation equations, and *a priori* coefficient sign, is contained in Table 2. The specification of models across time is largely consistent. The only difference is that a different non-wage income calculation is used in the SMA. Also, the loan commitment variable is unavailable from the SMA. A comparable explanatory variable specification is also used in the retirement age expectation models in Section 6.

Table 2. Labour force participation model explanatory variable specification

Explanatory Variable	Specification and values taken	*a priori* effect on labour force participation
University Qualifications - (DEGREE)	= 1 if university degree, 0 otherwise	Ambiguous
Trade Qualifications - (TRADE)	=1 if Trade qualification, 0 otherwise	Ambiguous
Other Qualifications - (OTH)	=1 if other diploma or certificate qualifications, 0 otherwise	Ambiguous
Individual's Non-Wage Income Quintile - (INDIVYQ)	=1, 2, 3, 4, 5 according to income quintile	Negative
Income Unit Non-Wage Income Quintile - (UNITYQ)	= 1, 2, 3, 4, 5 according to income quintile	Negative
Marital Status - (MARITAL)	=1 if married, 0 otherwise	Ambiguous
Dependents - (DEP)	=1 if dependents aged < 15 years, 0 otherwise	Positive
Home Ownership - (HOWN)	=1 if own home, 0 otherwise	Negative
Loans - (LOANS)	= 1 if individual paying off mortgage or other loan	Positive

This study is only concerned with the qualitative labour force participation decision rather than the quantitative choice of the number of hours worked. As such, qualitative response models, and in particular discrete choice models, are relevant, where the dependent variable takes whole number values 0, 1, 2, …. (Greene 2000). In the case of a binomial model

specification the dependent variable represents the choice between two options, namely employment (Y_B = 1) versus non-employment (Y_B = 0). A multinomial model specification represents the choice between a number of options. In the case of labour force participation, the choice is specified as between the three labour force states of NILF (Y_M = 0), unemployment (Y_M = 1) and employment (Y_M =2). For ease of estimation the logit estimation method was chosen.

The estimated model coefficients $\hat{b}$, measuring the influence of explanatory variables upon the log of odds ratio, have no immediate interpretation. However, the coefficient can be converted to calculate the marginal effect of an explanatory variable upon the probability of the event occurring. This marginal effect upon the individual's probability of employment does not increase linearly with *x*. That is, the magnitude of the marginal effect of a change in *x* for the probability of the event occurring depends on the initial starting position of *x*. Therefore, a base representative individual probability is first calculated for an individual using a chosen set of characteristics (*x*). The marginal effects of the change in *x* variables for the probability of the event occurring can then be calculated relative to this representative individual case. Model adequacy is assessed with likelihood ratio tests, pseudo R-square statistics and summary measures of the model's predictive success.

5. Binomial logit model results

The estimation results from the binomial logit models, conducted using the software package SPSS, are presented in Tables 3. and 4. Goodness of fit statistics, using both coefficient accuracy and overall predictive success criteria are reported. The estimated coefficient values ($\hat{b}$ **)** and their statistical significance (using Wald tests and the chi-squared distribution) are also reported. These coefficient estimates provide the marginal effects of changes in the explanatory variables upon the dependent variable, being the logit, or log of odds of employment (h>0) versus non-employment (h=0). As such, these coefficient estimates must be transformed to provide the probability of employment in the Pr column of

Tables 3 and 4. The control case or representative individual consists of the probability of employment for a married home owner, with no loans, no post secondary education or Trades qualifications, in the middle (3^{rd}) non-wage income quintiles. The probability of employment for the representative individual is reported next to the constant in Tables 3 and 4. The probability of employment is then recalculated separately for the above representative individual, but changing one of the explanatory variable values (generally setting the dummy value from zero to one, or vice versa). That is, the probability reported alongside DEP, DEGREE, TRADE, OTH, and LOAN coefficients provide the probability of employment for the representative individual but who now possesses dependents, university degree qualifications, Trades qualifications, other diploma or certificate qualifications, or loan commitments, respectively (setting the respective dummy variables from zero to one). The probability reported alongside the MARITAL and HOWN coefficients is for the representative individual now unmarried, or not a home owner, respectively (that is after setting MARITAL and HOWN dummies from one to zero, respectively). The probability alongside INDIVQ and UNITYQ is for the representative individual shifting up one non-wage income quintile (that is after setting INDIVYQ and UNITYQ from three to four, respectively). The marginal effect of that explanatory variable upon the probability of employment for the representative individual may then be calculated as the difference between the probability attached to the constant and that attached to the respective explanatory variable. For example, in 1973 the probability of employment for the representative individual described above is 0.842 or 84.2% for a male aged 55-59 years in Table 3. The same representative individual, except possessing other diploma or certificate qualifications (that is, setting all explanatory variables the same as the representative individual except setting OTH = 1), has a probability of employment of 0.962 or 96.4%. That is, the possession of these diploma or certificate qualifications increases the probability of employment for the married home owner with no dependents or loan commitments in the middle non-wage income quintile by around twelve percentage points.

Starting with the overall model goodness of fit, the likelihood ratio hypothesis test (LR) that all slope coefficients are equal to zero is rejected for all models at the one percent level of

significance (compared to critical chi-square distribution value with nine degrees of freedom and 1% level of significance of 21.666). The same log likelihood information ($\ln \hat{L}$ and $\ln \hat{L}_0$) is used for the calculation of the MacFadden's Pseudo R^2. While there is no natural interpretation for this pseudo R^2 between the values of zero and one as opposed to the conventional coefficient of determination,[14] they are observed to be relatively small and closer in value to zero than one. Furthermore, and although not directly comparable, it is observed that the pseudo R^2 appears to be relatively smaller for the 1973 equations.

Further goodness of fit information is provided by the summary of the model's predictive success rate.[15] All models allocate (predict) over 75% of individuals to their correct (actual) labour force state. However, this must be compared to the prediction success rate for the naïve prediction rule to assess the model's relative worth. In this regard, the 1973 equations performed quite poorly. The naïve prediction rule that all individuals in the sample are employed (that is, all $\hat{Y}_B = 1$) would achieve a prediction success rate of 87.948% for males aged 55-59 years and 78.082% for males aged 60-64 years in 1973. However, the estimated model's prediction success rate is only 87.622% and 75.685%, respectively. That is, the estimated model from 1973 achieves a lower prediction success rate than that achieved by the naïve prediction rule. In contrast, the predictive success rate from the 1985-86 and 1997-98 models are a vast improvement upon than that achieved by the naïve prediction rule. While the hypothesis that all explanatory variables provided no explanatory power was rejected in all models using the likelihood ratio test, it was evident that the accuracy of predictions from the 1973 models were quite poor, with a naïve prediction rule obtaining greater accuracy than that based upon model results.

With respect to the individual coefficients ($\hat{b}$) reported in Tables 3 and 4, only the individual's non-wage income (INDIVYQ) and market wage (human capital proxy) variable coefficients were statistically significant across models of different age groups and time periods. The effect of an increase in the individual's non-wage income quintile (INDIVYQ) reduced the log of odds of employment versus non-employment in all equations. In contrast,

the possession of formal education or training, representing potential market wages, generally had the effect of increasing the log of odds of employment, although this effect was not statistically significant across all models. The possession of university qualifications (DEGREE) increased the log of odds of employment for older males in both age groups in 1985-86 and 1997-98. However, its influence in 1973 was not statistically significant. Similarly, the possession of other diploma or certificate qualifications (OTH) significantly increased the log of odds of employment in three out of six equations. Notably, the possession of Trades qualifications (TRADE) had a statistically significant positive influence in only one model. Therefore, while the possession of tertiary education appears to exert a significant positive influence upon the log of odds of employment, the finding with respect to formal training is not so clear cut.

In comparison to the budget constraint variables, the effect of variables capturing the influence of observable characteristics upon the log of odds of employment were extremely inconsistent across age groups and time periods. That is, the coefficient sign attached to DEP, MARITAL, HOWN and LOAN differs across the different age group and time period models. Furthermore, the coefficients attached to these variables are rarely statistically significant. For example, the coefficients attached to marital status and loans, were only statistically significant in one equation each. The finding for MARITAL was especially surprising given previously identified empirical finding of unmarried males being more likely to be NILF and unemployed (for example, O'Brien 1999 and 2000a). This finding is either the reflection of a poorly specified model or that the previously observed significance of marital status was in fact proxying influences now statistically significant in the current study. The coefficient attached to the dependents variable (DEP) was statistically insignificant in all equations. However, this finding is not surprising given that relatively few older males in these age groups possess dependents (for example see Tables 5.15 and 5.16). In comparison, home ownership (HOWN) exerted a statistically significant influence in three equations. However, this home ownership coefficient displayed the opposite sign for males aged 60-64 years in 1985-86 versus 1997-98, hardly evoking confidence in the purported influence.

Of course, the estimated coefficients have greater intuitive appeal if expressed in terms of their marginal effect upon the individual's probability of employment. This is presented in the Pr columns of Tables 3 and 4. For males aged 55-59 years Table 5.4 shows that the representative individual[16] had a probability of employment of 84.2% in 1973, 63.2% in 1985-86 and 66.1% in 1997-98. In Table 4 the representative individual aged 60-64 years had a probability of employment of 74.7% in 1973, 47.3% in 1985-86 and 40.6% in 1997-98. These estimates appear plausible. That is, and as expected, the probability of the male aged 55-59 years being employed exceeds that for the male aged 60-64 years in each time period. Furthermore, the dramatic fall in probability of employment for the representative individual from 1973 to 1985-86 and subsequent low in 1997-98 seem especially plausible, given that that they approximate the aggregate labour force participation rate patterns (or more precisely the employment to population rate) in these respective years.

The representative individual depicted is one of many different individuals that may actually exist amongst the population. Tables 3 and 4 also show the probability of employment for the representative individual depicted above, but after altering one of the observable characteristics, market wage, or non-wage variables. A quick glance at Tables 3 and 4 indicates that the marginal effect of most explanatory variables upon the probability of the representative individual's employment is sizeable. However, with reference to the statistical significance of various coefficients in the logit equations from which the probability estimates are derived, we should really only have confidence in market wage (human capital) and non-wage variables, rather than the observable characteristics probability estimates. For example, the possession of university degree (DEGREE) qualifications increases the probability of the representative older male aged 55-59 years in 1985-86 being employed from 63.2% to 95.9%. Similarly, the possession of other diploma or certificate qualifications (OTH) increases the probability of employment to 81.1%. However, the possession of Trades qualifications only increases the probability of employment to 72.2% (and furthermore had a statistically insignificant effect on the log of odds of employment). In contrast, an increase in the representative individual's non-wage

income quintile from the third to fourth quintile reduces the probability of employment substantially to 34.5%. Other findings suggest that non-home ownership reduces the probability of employment to 42.9%, while loan commitments increase the probability of employment for the representative individual to 83.2%. Consistent with the observation of statistical significant coefficients, generally the market wages (human capital) and non-wage income variables had the greatest influence upon the representative individual's probability of employment.

Two main conclusions may be drawn from these binomial logit model results. First, the market wage (human capital) and non-wage income variables from the budget constraint were most important in explaining the labour force participation outcome, rather than the observable characteristics from the utility function. Consistent with overseas findings, this may mean that unobservable characteristics have a large role to play. Second, the battery of diagnostic tests revealed the important deficiencies in the model estimates. The models for 1973 all passed the basic likelihood ratio tests reported in overseas research, however, the naïve prediction rule provided greater accuracy than that based upon model results, meaning that some estimation results were not that fruitful. This casts doubt upon the findings from other models that have only presented rudimentary diagnostic results. However, the factors affecting the probability of labour force participation of those unemployed and NILF may have been restricted by the binomial model specification. The findings from the more general multinomial logit models are now reported.

Table 3. Binomial logit regression results of older male labour force participation – Males aged 55-59 years

	1973		1985-86		1997-98	
Variable	$\hat{b}$	**Pr**	$\hat{b}$	**Pr**	$\hat{b}$	**Pr**
MARITAL	1.184***	0.620	0.049	0.620	-0.357	0.736
DEP	0.738	0.918	0.590	0.756	0.555	0.773
DEGREE	-0.567	0.752	2.619***	0.959	1.630***	0.909
TRADE	0.344	0.883	0.415	0.722	0.544*	0.771
OTH	1.617***	0.964	0.920**	0.811	0.496	0.762
HOWN	-0.152	0.861	0.824**	0.429	0.584	0.521
LOANS			1.073**	0.834	0.433	0.750
INDIVYQ	-0.481***	0.767	-1.182***	0.345	-1.091***	0.396
UNITYQ	-0.037	0.837	-0.033	0.624	0.328**	0.730
Constant	2.122***	0.842	3.313***	0.632	2.731***	0.661
L0	-112.964		-305.285		-240.380	
L	-100.830		-199.588		-160.096	
LR	24.268		211.392		160.567	
MacFadden R^2	0.107		0.346		0.334	
Actual predicted %	87.622		84.848		83.168	
Naï ve Predicted %	87.948		69.293		71.782	

* statistically significant at the 10% level
** statistically significant at the 5% level
*** statistically significant at the 1% level

Table 4. Binomial logit regression results of older male labour force participation – Males aged 60-64 years

	1973		1985-86		1997-98	
Variable	$\hat{b}$	**Pr**	$\hat{b}$	**Pr**	$\hat{b}$	**Pr**
MARITAL	0.456	0.651	-0.157	0.512	-0.390	0.503
DEP	1.568	0.934	0.073	0.491	-0.029	0.399
DEGREE	7.415	1.000	0.981*	0.705	1.416***	0.738
TRADE	0.470	0.825	0.216	0.527	0.079	0.426
OTH	-0.182	0.711	1.130***	0.735	0.323	0.486
HOWN	0.407	0.663	0.586*	0.333	-0.632*	0.563
LOANS			0.526	0.603	0.620	0.560
INDIVYQ	-0.462***	0.650	-1.156***	0.220	-1.133***	0.181
UNITYQ	-0.094	0.729	0.064	0.489	0.235*	0.464
Constant	1.697***	0.747	2.741***	0.473	3.338***	0.406
L0	-153.553		-349.587		-256.643	
L	-140.277		-251.136		-176.039	
LR	26.551		196.901		161.207	
MacFadden R^2	0.086		0.282		0.314	
Actual predicted %	75.685		79.417		79.412	
Naï ve Predicted %	78.082		58.447		55.882	

* statistically significant at the 10% level
** statistically significant at the 5% level
*** statistically significant at the 1% level

6. Multinomial logit model results

The binary logit specification was driven by the orthodox theory of labour force participation distinguishing between those employed ($h > 0$) versus non-employed ($h = 0$), with unemployed and NILF depicted as operationally equivalent states. However, stylised facts and research issues in Chapter 2 were presented in terms of distinguishing the unemployed from those NILF using the labour force framework. Therefore, an interesting empirical proposition to address is whether we can statistically distinguish between the two non-employment states of unemployment and NILF in an econometric model. A multinomial specification allows the calculation of probabilities of employment, unemployment and NILF for an individual with given characteristics within one model. This may provide an indirect test of whether the binomial specification was justified on theoretical grounds or was overly restrictive.

Two equations for each age group and year were estimated for the multinomial model. The first equation in Tables 5 and 6 reports the estimated coefficients for the log of odds of employment versus NILF, with the second equation relating to the log of odds of unemployment versus NILF. The same goodness of fit measures are reported as for the binomial model. The probabilities of employment and unemployment are reported in Tables 7 and 8 alongside each explanatory variable based upon the same representative individual as used in the binomial model.

Starting with the goodness of fit tests, all likelihood ratio hypothesis tests that explanatory variable coefficients are jointly equal to zero are rejected, meaning that the models adequately fit the data. Pseudo R^2 values are closer to zero than one, as with the binomial models. The prediction rates are slightly lower than the binomial model. The naïve prediction rule is again more accurate in 1973 model than that based upon the model estimates. In summary, although most diagnostics are not directly comparable, there has been no apparent improvement in model fit from the multinomial specification.

With reference to coefficient estimates and statistical significance in Tables 5 and 6 it is clear that the main explanatory variables affecting the log of odds of employment versus NILF are similar to those from the binomial specification of employment versus non-employment. For example, for males aged 55-59 years in 1985-86, the same coefficients are statistically significant except for the loan commitments coefficient.

The main finding is that few characteristics differentiate the older male unemployed from those NILF. Statistically significant coefficients are rare in the equation estimating the effect of explanatory variables upon the log of odds of unemployment versus NILF. The only exception is that those possessing a university degree are less likely to be unemployed in three out of six equations.

These findings are also reflected in the estimated probabilities of employment and unemployment reported in Tables 7 and 8. Similar probabilities of employment to the binomial regression are reported. For example, in 1985-86 the representative 55-59 year old male had a probability of employment of 64.7% (cf 63.2% binomial regression), increasing to 95.8% with a university degree (cf 95.9%), 73.6% with Trades qualifications (cf 72.2%), 82.5% for other diploma or certificate qualifications (cf 81.1%), and drops to 35.8% with an increase in non-wage income quintile (cf 34.5%). However, the probability of the representative 55-59 year old in 1985-86 of being unemployed is 4.5%, decreasing to 1.6% for a university degree holder, and increasing to 8.4% with an increase in non-wage quintile (although this effect may actually reflect the receipt of unemployment benefits). Notably, the representative unemployed 60-64 year old in 1973 was indistinguishable from the NILF individual, displaying a zero probability. Presumably, this is partly because there were so few unemployed older males in that year. In general, the probability of unemployment for the older male was less than 5%.

In summary, the move from the binomial logit to the multinomial logit model did not offer better insights into older males' labour force participation behaviour. Similar findings for employment versus NILF were obtained, as were available from the binomial specification

of employment versus non-employment. With reference to the statistical significance of coefficients and probability estimates, the older male unemployed and NILF were largely indistinguishable. This has two implications. First, there are few differences between the older unemployed and NILF. This implies that essentially unemployment and NILF are equivalent states for older males. Second, the findings suggest that the binomial model specification was not overly restrictive.

Table 5 Multinomial logit regression results of older male labour force participation – Males aged 55-59 years

	1973		1985-86		1997-98	
Variable	**E/NILF** $\hat{b}$	**U/NILF** $\hat{b}$	**E/NILF** $\hat{b}$	**U/NILF** $\hat{b}$	**E/NILF** $\hat{b}$	**U/NILF** $\hat{b}$
MARITAL	1.186**	-0.058	0.406	1.723***	-0.398	-0.191
DEP	0.685	-14.857	0.606	-0.018	0.649	0.621
DEGREE	-0.710	-16.461	2.892***	1.493	1.499**	-15.616***
TRADE	0.584	1.776	0.502	0.512	0.753	1.123*
OTH	1.884***	2.087	0.966**	0.166	0.565**	0.434
HOWN	-0.293	-1.628	0.650*	-0.694	0.665	0.536
LOANS			0.744	-1.995***	0.748	1.567*
INDIVYQ	-0.483***	0.060	-1.186***	0.025***	-1.097***	0.001
UNITYQ	-0.011	0.194	-0.166	-0.831	0.311**	-0.118
Constant	2.270***	-2.118	3.741***	-0.553	2.843***	-2.326***
L0	-80.531		-296.584		-243.769	
L	-66.114		-178.630		-158.734	
LR	28.834		235.908		170.071	
MacFadden R^2	0.179		0.398		0.349	
Actual predicted %	87.622		81.818		79.703	
Naï ve Predicted %	87.948		69.293		71.782	

* statistically significant at the 10% level
** statistically significant at the 5% level
*** statistically significant at the 1% level

Table 6 Multinomial logit regression results of older male labour force participation – Males aged 60-64 years

	1973		1985-86		1997-98	
Variable	**E/NILF** $\hat{b}$	**U/NILF** $\hat{b}$	**E/NILF** $\hat{b}$	**U/NILF** $\hat{b}$	**E/NILF** $\hat{b}$	**U/NILF** $\hat{b}$
MARITAL	0.165	-54.405	0.112	0.456	-0.459	0.155
DEP	1.577	38.849	-0.177	0.897	-0.111	-0.252
DEGREE	17.358***	40.989	0.959*	-15.037***	1.421***	-15.078***
TRADE	0.586	41.678	0.317	1.151**	0.143	0.620
OTH	-0.260	-43.955	1.221***	1.251	0.617	1.675**
HOWN	0.266	-40.342	0.642*	0.783	-0.795**	-1.279
LOANS			0.450	-15.618	0.590	-0.333
INDIVYQ	-0.438***	13.696	-1.190***	-0.382	-1.244***	-0.845
UNITYQ	-0.059	7.302	0.031	-0.386	0.318**	0.527*
Constant	2.025***	-43.671	2.932***	-1.557	3.742***	-0.802***
L0	-107.540		-297.796		-241.308	
L	-81.996		-190.720		-149.689	
LR	51.087		214.153		183.238	
MacFadden R^2	0.238		0.360		0.380	
Actual predicted %	76.712		77.864		78.342	
Naï ve Predicted %	78.082		55.146		52.406	

* statistically significant at the 10% level
** statistically significant at the 5% level
*** statistically significant at the 1% level

Table 7 Probability of employment and unemployment – Males aged 55-59 years

	1973		1985-86		1997-98	
Variable	**Pr (E)**	**Pr (U)**	**Pr (E)**	**Pr (U)**	**Pr (E)**	**Pr (U)**
MARITAL	0.618	0.013	0.577	0.011	0.739	0.027
DEP	0.916	0.000	0.771	0.029	0.775	0.034
DEGREE	0.729	0.000	0.958	0.016	0.905	0.000
TRADE	0.892	0.017	0.736	0.052	0.776	0.051
OTH	0.966	0.007	0.825	0.026	0.764	0.031
HOWN	0.863	0.019	0.459	0.122	0.508	0.026
LOANS			0.813	0.004	0.754	0.078
INDIVYQ	0.766	0.008	0.358	0.084	0.392	0.054
UNITYQ	0.839	0.006	0.626	0.022	0.727	0.022
Constant	0.842	0.005	0.647	0.045	0.659	0.030

Table 8 Probability of employment and unemployment – Males aged 60-64 years

	1973		1985-86		1997-98	
Variable	**Pr (E)**	**Pr (U)**	**Pr (E)**	**Pr (U)**	**Pr (E)**	**Pr (U)**
MARITAL	0.715	0.000	0.464	0.024	0.530	0.022
DEP	0.935	0.000	0.419	0.088	0.391	0.025
DEGREE	1.000	0.000	0.725	0.000	0.756	0.000
TRADE	0.842	0.000	0.531	0.087	0.439	0.053
OTH	0.695	0.000	0.733	0.054	0.516	0.111
HOWN	0.694	0.000	0.341	0.021	0.580	0.070
LOANS			0.614	0.000	0.565	0.017
INDIVYQ	0.656	0.000	0.227	0.036	0.174	0.019
UNITYQ	0.736	0.000	0.499	0.024	0.484	0.045
Constant	0.747	0.000	0.486	0.035	0.415	0.031

7. An analysis of retirement expectations

An aspect of orthodox theory of labour force participation emphasised in the lifecycle model under conditions of certainty is the role of retirement expectations and planning. However, the analysis of retirement expectations remains a relatively neglected topic in Australian and overseas literature (an exception to this in overseas literature is Holtmann *et al*. 1994). With regard to Australian empirical studies, Miller (1983) found that retirement expectations taken in 1973 ultimately predicted the bulk of actual labour force participation rates of older males in the proceeding decade. However, while Miller stated that expected receipt of the ASP and increasing levels of income and wealth over time explained the earlier retirement expectations and subsequent older male labour force participation rates,[17] these propositions were never directly analysed. Borland (1995) later found that retirement expectations were a poor predictor of actual behaviour.

Only two sources of retirement expectation data are available in Australia. Namely the SMA which Miller (1983) used, and the ABS Retirement and Retirement Intentions Survey used by Borland (1995). However, as mentioned previously, the ABS data is not available to the public in CURF format. As a result, we are left to analyse the SMA which is dated and doesn't allow the more recent inaccuracy of projections issue to be directly addressed. However, the use of the SMA allows an indirect test of two important issues put forth in Miller's work. First, the role of increasing wealth and income, and expected receipt of ASP, purported by Miller (1983) to be the key to earlier retirement expectations. Second, a comparison is made between retirement expectations and actual labour force participation behaviour.

In this section, two models of retirement expectations are presented. In the first model, the expected age of retirement (in years) is the dependent variable. This is a continuous variable and therefore a linear regression model is specified and estimated. Coefficients measure the marginal effect, in years, of explanatory variables upon the expected age of retirement. Second, a binomial logit model is specified, with the dependent variable taking the value of 1 if the individual expects to retire prior to 65, and zero otherwise. Estimated coefficients

can be converted to give the marginal effect of explanatory variables upon the probability of expected retirement prior to 65 years for individuals with various characteristics.

The specification of explanatory variables is similar to the above binomial and multinomial logit models of actual labour force participation. That is, financial variables representing the budget constraint and observable characteristics from the utility function in a reduced form specification. In contrast to the labour force participation models presented above, all age groups are included in the model sample, rather than being restricted to the 55-64 year old age group. This is simply because we are interested in retirement expectations rather than actual behaviour. Age group (AGE) is thus included as an explanatory variable to capture the potential influence of different generations' retirement expectations. The sample is restricted to those currently employed and thus excludes those who have already left the labour force. This means that actual market wage income quintile (INCQ) can be used rather than human capital proxies. It was contended by Miller (1983) that those with higher income anticipate earlier retirement than lower income groups because of an intertemporal income effect. The role of increasing income over decades emphasised by Miller (1983) cannot be directly captured in the model. However, the influence of income level upon retirement expectations in a cross section can be modelled. A variable is included to capture whether the individual was a war veteran (VETERAN). This should capture the purported influence of expected receipt of the ASP at age 60.

Although most coefficients in the linear model in Table 9 are statistically significant, and display the expected sign, they typically reveal only a small influence (less than two years) upon the expected age of retirement. The representative individual[18] had an expected age of retirement of 63.253 years. An increase in the individual's age group increases this expected retirement age by 1.544 years to 64.796 years. This marginal effect could give credence to Miller's claim that successive generations have a lower expected retirement age. However, this finding may also be clouded by the fact that some older males who may have expected earlier retirement had already left the labour force and were therefore not part of the estimation sample. Similarly, those married or with dependents expect to retire 1.306 years

and 0.293 years later, respectively. In contrast, being a veteran, and thus anticipating the ASP at 60 years of age, rather than 65 years for the Age pension, reduces the expected retirement age by 1.197 years to 62.056 years. While statistically significant, the marginal effect is hardly large, considering these veterans had an expected receipt of social security five years earlier than other groups. An increase in market wage income quintile decreases the age of expected retirement by 0.437 years to 62.816 years. This finding is also consistent with Miller's contentions. However, the small quantitative influence revealed is unlikely to have explained the dramatic and sudden drop in older male labour force participation rates experienced over the 1970's to 1980's. Perhaps of greater importance, the income finding is in contrast to the observation from actual behaviour in labour force participation models estimated previously which indicated that those with higher market wages actually cease working at later ages compared to lower income earners.

The findings presented in Table 10 show that the binomial logit model satisfies the likelihood ratio test (that is, the null hypothesis of zero explanatory power is rejected). The model's predictive success rate is relatively low at 66.392%, although this represents a fourteen percentage point improvement upon the naïve prediction rate. All coefficients apart from HOWN are statistically significant at the 1% level of significance. As expected, coefficients display the opposite sign to those from the linear model because of the dependent variable specification being the logit of early retirement expectations (1) versus retirement at 65 years or later (0). The representative individual displays a relatively low probability of having an early retirement expectation of 32.1%. An increase in age group reduces this probability to 20.1%. Being unmarried or a veteran increases the probability of early retirement expectation to 45.1% and 50.1% respectively. Therefore, while being a veteran has a large impact on the probability of having early retirement expectations, the previous linear equation suggested this actually had only a small influence upon expected years of early retirement. In comparison, market wage income had a very small influence upon the probability of early retirement expectations. The marginal effects of changes to other explanatory variables for the probability of early retirement expectations were less than five percentage points.

In summary, the analysis of retirement expectations seemed to generate more mysteries than insights. First, the expectation of ASP receipt, and higher income, which were suggested by Miller (1983) as the driving forces behind earlier retirement expectations and thus older male labour force participation trends, were statistically significant but had a relatively small quantitative influence on years of early retirement expectations or the probability of early retirement expectation. Second, with respect to market wage income, there is a divide between those who expect to retire earlier, and those who actually do. This issue was not addressed by Miller (1983) in his research. According to the models presented above, higher income individuals expect to retire earlier. However, it was clear from both binomial and multinomial logit equations in Sections 5 and 6 that those with lower actual or potential labour market earnings were in fact the individuals that had left the labour force earlier. Therefore, while the wealthier expect to leave the labour force earlier, it is the less well off that actually do exit earlier. This finding suggests that lower wage earners' labour force participation trends may be driven more by unanticipated events. This accords with Phelan and Rust's (1991) U.S. analysis which suggested less accurate prediction of retirement timing for lower wage workers. This is also an important finding contradicting Merrilee's (1986) claimed consensus on the lifecycle model under conditions of certainty driving the long-term trend in older male labour force participation rates.

Table 9 Expected retirement age linear equation results

Variable	Coefficient	Expected age of retirement
AGE	1.544 (16.156)***	64.796
VETERAN	-1.197 (-4.707)***	62.056
DEP	0.293 (1.795)*	63.546
MARITAL	1.306 (3.245)***	61.947
HOWN	-0.159 (-0.509)	63.094
INCQ	-0.437 (-5.253)***	62.816
UNITYQ	-0.128 (-1.931)*	63.125
Constant	58.755 (98.017)***	63.253
Adjusted R^2	0.172	
Durbin-Watson	1.979	
s.e. of estimate	4.40	

* statistically significant at the 10% level
** statistically significant at the 5% level
*** statistically significant at the 1% level

Table 10 Binomial logit regression results for expected early retirement

Variable	$\hat{b}$	Pr(early)
AGE	-0.633***	0.201
VETERAN	0.754***	0.501
DEP	-0.215***	0.276
MARITAL	-0.553***	0.451
HOWN	0.158	0.357
INCQ	0.279***	0.385
UNITYQ	0.100***	0.344
Constant	0.766***	0.321
L0	-1091.494	
L	-1918.900	
LR	264.088	
MacFadden R^2	0.431	
Actual predicted %	66.392	
Naï ve Predicted %	52.251	

* statistically significant at the 10% level
** statistically significant at the 5% level
*** statistically significant at the 1% level

8. Macro inferences

This section presents a rudimentary analysis of whether changes in variables at the aggregate level suggested by orthodox theory and micro econometric model results can help account for older male labour force participation (or more precisely, employment to population) rates over time. For example, has there been an aggregate decline in older male human capital, a decline in labour market wages, an increase in wealth, and/or an increase in social security pension value and subsequent stabilisation at these new levels which, *ceteris*

paribus, may explain the initial fall in the older male labour force participation rate and sustained low level? A more technical analysis has not been conducted for a number of reasons. First, micro based orthodox theory of labour force participation doesn't directly predict macro outcomes without restrictive assumptions. Therefore, it is strictly an empirical proposition to see if macro outcomes are consistent with micro predictions. Second, inconsistent and inaccurate estimates of observable characteristics coefficients in logit equations means that a technical microsimulation using different population compositions over time is out of the question. However, the analysis of the change in human capital composition of the older male population may prove fruitful. Unit record data from the SMA and the SIHCs had population weights applied to provide population estimates of older male labour force status and human capital composition in Tables 11 and 12. Aggregate observable characteristics for older males are also reported but are not directly analysed because of their statistical insignificance in the estimated equations.

An important finding from the micro-based orthodox models of labour force participation is that those older males with higher potential market wages, as proxied by DEGREE and OTH variables, are more likely stay employed at older ages. Marrying micro results with macro trends implies that, *ceteris paribus*, lower older male labour force participation rates may be explained by a lower percentage of older males possessing these human capital qualities over time. It is clear in Tables 11 and 12 that this is not quite the case. There is generally a marginal increase in the proportion of older males possessing degree qualifications over time but a greater decrease in the proportion possessing other qualifications. However, the small net decrease in the proportion of the older population possessing tertiary qualifications would be unable to account for the larger decrease in the older male labour force participation rate over time without other counter veiling forces.

Table 11 Labour force status and population composition males 55-59 (% of age group)

Variable	1973	1985-86	1997-98
Employed	86.7	69.9	72.4
Unemployed	0.7	5.0	5.1
NILF	12.6	25.1	22.5
Married	89.0	82.2	83.0
Dependents	4.5	14.8	18.6
Home owners	87.4	61.1	81.8
Degree	4.8	5.8	7.6
Trades	17.8	24.4	28.5
Other qual	19.2	9.4	12.5

Source: SMA, SIHC 1985-86 and 1997-98

Table 12 Labour force status and population composition males 60-64 (% of age population)

Variable	1973	1985-86	1997-98
Employed	73.7	42.0	44.6
Unemployed		3.9	3.5
NILF	0.6	54.1	51.9
Married	25.6 86.7	82.1	78.8
Dependents	0.9	6.9	7.0
Home owners	84.3	65.8	90.2
Degree	4.9	5.4	3.5
Trades	16.0	25.2	30.2
Other qual	17.3	9.6	10.2

Source: SMA, SIHC 1985-86 and 1997-98

Human capital was proxy for labour market wages in the previous modelling. Figure 2 shows that real average weekly earnings (AWE) has exhibited a steady increase from the mid 1960's to mid 1980's before a fall during the ACCORD years (mid 1980's to mid 1990's), increasing again from the mid 1990's. In general, real labour market earnings have increased rather than fallen. In particular, earnings were increasing rapidly during the 1970's when older male labour force participation rates fell, contrary to the micro prediction of a decline in aggregate or average labour market wages.

Figure 2 Real Average Weekly Total Earnings ($1989-90) Males, Australia 1966-1999

Source: ABS (1999c, 1999d)

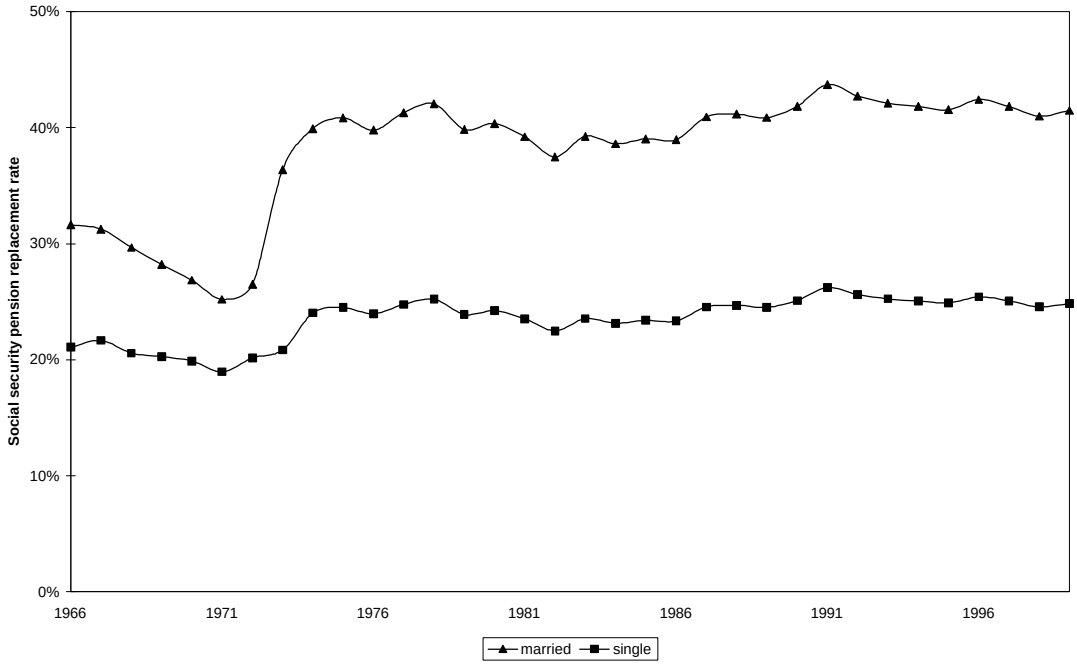
Hughes (1984) suggested that the long-term accumulation of labour force earnings and the lifecycle model of labour supply were dominant in explaining older male labour force participation rates. The twenty year average of annual real AWE changes, representing growth in labour market earnings accumulation, is presented in Figure 3. It is clear that earnings accumulation growth had occurred up until the late 1970's, potentially allowing earlier labour force exit than prior cohorts. This era was clearly over from the mid 1980's with a long-term decline in accumulation growth apparent after this time until the late 1990's. Therefore, while the lifecycle model and wealth accumulation may have allowed an explanation of the initial decline in older male labour force participation rates in the 1970's, it is unlikely that the same explanation can be used to account for the continued decline and sustained lower older male labour force participation rates thereafter.

Figure 3 Twenty Year Average of Real Average Total Earnings Increases ($1989-90) Males, Australia 1966 to 1999

Source: ABS (1999c, 1999d)

Perhaps a greater insight into the relative financial attractiveness of non-wage income and NILF status is provided by social security pension replacement rates. That is, the maximum value of social security pensions as a ratio of AWE. Figure 4 shows that there was a rapid increase in the value of social security pensions in 1973, especially for married couples. However, social security pension replacement rates have remained constant after the mid 1970's at around 25% and 40% for single persons and married couples respectively. Therefore, *ceteris paribus*, the increase in value of non-wage income may explain the decrease in older male labour force participation rates and increase in social security pension receipt in a limited period in the 1970's only.

Figure 4 Social security pension replacement rates (%) 1966 to 1999



Source: ABS (1999c), DFaCS (2000b)

In summary, aggregate changes in budget constraint variables or population composition do not appear to exhibit trends consistent with older male labour force participation rate trends over time. This finding has major implications. Primarily, a problem of aggregation leads to a fallacy of composition in reasoning, whereby what holds at the micro level does not hold true at the macro level. According to the findings from micro models, individuals with higher actual and potential market wages are likely to have higher employment and labour force attachment, while an increase in non-wage income value increases the probability of an individual's non-employment status. However, *ceteris paribus*, trends in older male labour force participation rates do not appear to reflect the trends in older male population human capital composition, market wages, long-term market wage accumulation, or social security pension replacement rates.

9. Conclusions

The research findings presented in this paper emphasised the role of human capital and potential market wages for an individual's labour force participation. However, overall the findings from unit record analysis of orthodox theory propositions results were disappointing. Rather than confirming theoretical propositions, the preceding analysis provided a number of anomalies. First, even though all models satisfied the likelihood ratio tests, the predictive success of the models compared to the naïve prediction rule was mixed. For example, better prediction results were available with the naïve rule in 1973 than from the estimated model. This finding emphasises the importance of presenting a battery of diagnostic tests to assess model adequacy. Therefore, one should caution some of the earlier logit based research that did not present evidence of satisfactory diagnostics performance. Second, the coefficients attached to observable characteristics variables, representing tastes and preferences for work and leisure, were rarely statistically significant or of a consistent sign across age groups or time periods. This may mean that unobservable tastes dominate the labour force participation decision. Although this is accommodated for in the utility function and reduced form econometric model specification (e), it provides little guidance for empirical estimation and prediction. Third, the general multinomial logit specification offered no improvement on the binomial logit. The model was largely unable to distinguish the factors influencing the older male unemployed from those NILF. Fourth, while it was shown that those with higher potential wages were likely to remain in the labour force at older ages compared to lower wage earners, higher wage earners had greater expectations of earlier retirement than lower earners. This shows that expectations do not necessarily match actual labour force participation behaviour, and that lower wage earners may be more vulnerable to unanticipated labour force exit. Finally, micro findings were not adequately translated to the macro level, being unable to directly offer inferences for aggregate changes in older male labour force participation rates over time. This problem of aggregation results in a fallacy of composition in reasoning. This means that the micro models presented have limited use for explaining the stylised facts.

Rather than arguing that orthodox theory of labour force participation simply cannot explain older male labour force trends, the findings perhaps imply that there may be a large role for unanticipated events and unobservable tastes and preferences. Unfortunately, without a rich longitudinal dataset, these forces are difficult to predict and incorporate into an empirically tractable lifecycle model under uncertainty.

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[1] See Merrilees (1986) for a summary of the research findings from this era of research.
[2] A non-technical (and non-critical) summary of U.S. retirement literature from the 1970's to the early 1990's was provided by Leonesio (1996).
[3] See Pencavel (1986) for a thorough outline of these issues.
[4] Again, see Pencavel (1986).
[5] Again, see Leonesio (1996) for a summary of U.S. retirement literature from the 1970's to the early 1990's.
[6] Leonesio (1996: 13-14) says "In the reduced-form approach, economic theory *suggests* which explanatory factors are likely to influence labor supply or retirement and might also provide *some indication* about the type and form of statistical model to select. The resulting statistical model essentially confirms and measures the influence of the explanatory factors *thought to be associated* with the outcome. The reduced-form model is usually neither directly nor uniquely linked to the researcher's theoretical model."
[7] Leonesio (1996: 14) says "The defining nature of a structural model is that the underlying preferences believed to have generated the observed behaviour are specified and estimated. The advantage of the structural modeling approach is that once a mathematical representation of preferences has been determined, it is then possible to predict the response to a considerably larger variety of changes in the individual's opportunities than in the case of reduced-form models. That is, once individual preferences (that is, the utility function) are known, then it is possible to predict how behaviour will respond to changed opportunities. It is only through developing increasingly detailed and realistic structural models that one can determine the influence of plausible factors such as the specific features of the Social Security system, the structure of private pension plans, and so forth."
[8] Gruber and Wise (1998: 162) say "It seems evident that if the trend to early retirement is to be reversed, as will almost surely be dictated by demographic trends, changing the provision of social security programs that induce early retirement will play a role."
[9] The Australian longitudinal survey, available through the SSDA, started in 1985 (wave 1), sampling youth aged 15 to 24 years in September 1984, and who were unemployed and registered with the CES for at least 3 months. See www.anu.ssda.edu.au
[10] SEUP, available from the ABS, is targeted at those aged 59 years or under.
[11] Because of confidentiality agreements, this data cannot be displayed in the Data Appendix without ABS or SSDA approval. The author is happy to make arrangements for data access if required.
[12] Experience would also be a relevant human capital variable to augment this specification. However, econometric models are already classified by age group, while the exact age in years is not known for those in the 1985-86 SIHC. Furthermore, years employed, or employment duration is not known.
[13] Data to calculate LOANS was only available in the SIHC, not SMA.
[14] The conventional coefficient of determination, R^2, provides the proportion of the variation in the dependent variable that is explained by variation in the explanatory variables.
[15] That is, the percentage of individuals correctly predicted to be in their actual labour force state.
[16] That is, an individual who is a married home owner with no dependents, no loan commitments, or no post school education or training, in the middle non-wages income quintile.
[17] Miller (1983: 11) says "It is conjectured, therefore, that the changing retirement plans are largely reflections of change in individual preferences which derive from the greater life-time income/wealth of today's 60-64 year old cohort vis-à-vis earlier 60-64 year old cohorts."
[18] That is, aged 45-54 years, married, with no dependents, in the middle individual market wage income and middle income-unit non-wage income quintile, who is not a home owner, and not a veteran, is reported alongside the constant