New England Public Policy Center
at the Federal Reserve Bank of Boston

Working Paper

07-4

The Labor Market for Direct Care Workers

by Reagan Baughman and Kristin Smith
University of New Hampshire

Abstract

As the baby boom cohort nears retirement age, the question of how to provide necessary health care and personal services to a growing elderly population has become a looming policy problem. Beginning in 2020, the number of Americans over the age of 65 will surpass the number of primary providers of formal and informal long-term care (women between the ages of 20 and 44). Perceptions of shortage and very high turnover in today's direct care labor market are compounding the potential problem.

This paper provides an overview of the labor market for direct care workers in the United States, including comprehensive empirical analyses of wage determination and labor supply, using panel data from the 1996 and 2001 Surveys of Income and Program Participation (SIPP). The paper describes the ways in which public policy is expected to affect the direct care labor market and empirically analyzes the wages, health insurance coverage, and employment duration of direct care workers. The authors find both that wages of direct care workers are quite low, with median starting wage of $7.96, and that spells of employment with a given employer are short, averaging just under 10 months.

The authors gratefully acknowledge support from the Carsey Institute at the University of New Hampshire, the W.E. Upjohn Institute for Employment Research, and the New England Public Policy Center at the Federal Reserve Bank of Boston. Alicia Sasser, Christopher Foote, Robert Triest, Christine Bishop and other participants in the Federal Reserve Bank of Boston research seminar and the New England Study Group provided helpful suggestions. All errors are our own.

FEDERAL RESERVE BANK OF BOSTON™

1. Introduction

As the baby boom cohort nears retirement age, the question of how to provide necessary health care and personal services to a growing elderly population has become a looming policy problem in most western countries. Although a relatively small fraction of individuals over the age of 65 in the United States reside in nursing care facilities, almost 16 percent do receive some form of long-term care and the need for care increases dramatically with age (Green Book 2004). Given that much of the need for long-term care takes the form of assistance with activities of daily living, such as eating and grooming, the aging baby boom population is likely to put more strain on the market for direct care workers, such as nursing and home health aides,[1] than it will on other markets in the health care sector.

The impact of increased demand for all types of nursing services, and the pressure on the nursing labor market, is illustrated by the population projections in Figure 1. The number of Americans over the age of 65 will surpass the number of women between the ages of 20 and 44 (the primary providers of formal and informal long-term care), beginning in 2020. By 2050, the ratio of 20-to-44 year old women to the over-65 population, which is currently about 1.5, will decrease to approximately 0.67. Further, the gap between supply and demand for care could be even more dramatic than the figure would indicate, given that the entire nursing workforce has been aging over time as nursing careers (at all skill levels) have become less popular with young women (GAO 2001).

Compounding the potential problem is that there are at least perceptions of shortage and very high turnover in today's direct care labor market. Although groups including the Institute of Medicine have called for responses to these workforce problems, the GAO (2001) notes that "comprehensive data are lacking on the nature and extent of the shortage" and suggests that more work needs to be done on shortage and turnover, particularly for nursing aides. The work that has been done on this topic, usually at the state or nursing home level, suggests that serious problems do exist in the direct care labor market. In a recent national survey (PHI 2002), 37 of 43 states surveyed reported what they considered to be serious shortages of direct care workers. Additionally, state-level studies of turnover among care workers report annual rates that range from 25 percent to well over 100 percent

[1] In this paper the term "direct care worker" refers to paraprofessional health occupations, such as nursing aides, home health aides and personal care attendants. Descriptions of each of these occupations are provided in Section 2.

(Wright 2005). High levels of turnover have been shown to adversely affect patient outcomes in nursing home settings (Barry et al. 2005).

Unfilled positions and high turnover rates may not be surprising, given that median wages in 2006 for direct care occupations ranged from $8.54 per hour for personal and home care attendants to $10.67 for nursing aides, orderlies and attendants (BLS 2006). Low levels of required education and training are likely a partial explanation for these low wages. However, wages that remain low even in the face of excess demand are better explained by the fact that the public health insurance program, Medicaid, is the primary payer in the market for long-term care for the elderly. Because of economic and political constraints on public spending at federal and state levels, Medicaid reimbursement rates to long-term care providers are usually well below what private insurance companies pay. The implication for the direct care worker labor market is that wages may not necessarily be able to adjust in the face of changing demand conditions. Borjas (1983) documents a similar effect of Medicaid nursing home reimbursement rates on R.N. wages.

This paper provides an overview of the labor market for direct care workers in the United States, including comprehensive empirical analyses of wage determination and labor supply, using panel data from the 1996 and 2001 Surveys of Income and Program Participation (SIPP). The SIPP panels (four years in 1996; three years in 2001) provide a rich source of month-level data on more than 3,600 employment spells in direct care occupations. Using this month-level data allows us to capture employment duration very accurately in an occupation where turnover is high and job spells would therefore be expected to be short. In fact, the mean duration of spells with the same direct care employer in these data is only 9.7 months and the median spell length is five months; therefore, month-level data is critical for an accurate analysis of employment transitions.

The paper proceeds as follows. The next section provides a brief overview of each of the direct care occupations, including responsibilities, training requirements, and places of employment. Section 3 reviews relevant literature on nursing labor supply and Section 4 describes the ways in which public policy is expected to affect the direct care labor market. Section 5 describes the SIPP data used for analysis in more detail and is followed by sections devoted to empirical analyses of wages (6), health insurance coverage (7), and employment duration (8). Section 9 concludes.

2. Overview of the direct care labor market

In 2006, according to the Bureau of Labor Statistics, 2.7 million workers worked in the direct care occupations, with 51 percent working as nursing aides, orderlies or attendants, 28 percent working as home health aides, and the remaining 21 percent working as personal or home care aides. More than 90 percent of these workers were employed in the health care sector; the analysis that follows covers health care sector workers only.[2]

Nursing aides are predominantly Certified Nursing Assistants (C.N.A.s) and Licensed Nursing Assistants (L.N.A.s), the two occupations with the most stringent licensing standards. C.N.A.s are required by the federal government to have 75 hours of classroom training and 16 hours of practical training.[3] As Appendix Table 1 shows, 26 states require additional training, with minimum classroom hours ranging as high as 175 hours. Figure 2 breaks down employment for nursing aides and the occupations by industry. The two largest employment sectors for nursing aides, orderlies and attendants are nursing care facilities (47 percent) and hospitals (36 percent). The responsibilities of nursing aides in the two sectors are similar and include moving patients to and from bed, bathing, changing sheets, feeding and generally monitoring physical and mental health. An additional, important responsibility of aides in nursing homes is to monitor and ensure the safety of patients with dementia, who may either be restrained or under close nursing supervision.

Home health aides perform tasks that are very similar to those of aides in nursing homes, but in a more diverse set of workplaces. Despite the name, only 42 percent of home health aides actually work in the home health sector (Figure 2). Nearly 25 percent work in residential facilities for the mentally ill or disabled, 19 percent work in community care for the elderly, and 10 percent work in nursing homes. Home health aide is projected to be the single fastest-growing occupation in the United States between 2004 and 2014, with a 56 percent increase in total employment over the ten-year period.

Personal care or home care aide is the tenth fastest growing occupation over the same period, with a projected employment growth rate of 41 percent between 2004 and 2014 (BLS 2004). Personal and home care attendants are rarely responsible for the more

[2] The health care sector definition includes personal and health care services performed in a patient's home. It excludes the small number of individuals reporting direct care occupations in other industries, such as manufacturing.
[3] This compares to two to four years for Registered Nurses (R.N.s) and 12 to 18 months for Licensed Practical Nurses (L.P.N.s).

"medical" tasks sometimes performed by nursing and home health aides (i.e., monitoring vital signs or assisting with medication) but do assist patients with activities of daily living (i.e., bathing, grooming, dressing, eating, moving around) in the same way as nursing and home health aides. Personal and home care attendants are also likely to do some shopping, cleaning, laundry or other household work. The large majority (71 percent) of personal and home care attendants work in the home health sector, with the next largest sectors being mental health and other residential facilities (17 percent) and community care for the elderly (7 percent).

3. Existing models of nursing labor supply

While there are very few papers in the literature that analyze the employment behavior of direct care workers from an economic perspective, there is a larger literature that focuses on the labor market for Registered Nurses (R.N.s). Although R.N.s are a higher skilled workforce, with two-to-four years of required training, this literature suggests a number of relevant points to consider in an analysis of direct care labor supply. The summary below is separated into wage and labor supply analyses.

A. Wages

There are a number of papers in the R.N. workforce literature that attempt to explain why shortages of nurses occur in the United States (as opposed to European countries with nationalized health systems), where wages are seemingly set by market forces and would be expected to rise to meet demand in a competitive environment. The most common explanation is that hospitals enjoy monopsony power as the predominant employer of R.N.s in most geographic areas. As Figure 3 illustrates, when a single employer has market power such that it faces an upward sloping labor supply curve $w(L)$, the marginal cost of hiring an additional worker will be higher than the wage, as long as the employer is constrained to offer all employees the same wage. This produces employment level L_m and wage w_m, levels that are below the competitive levels (L_c and w_c), as well as a shortage of workers willing to work at w_m that is equal to (L^*_m-L_m). As is shown by Boal and Ramson (1997) the degree of monopsony (and the size of the wage and employment gaps) is an inverse function of elasticity of labor supply facing the firm.

There is not a consensus in the literature as to whether monopsony is actually responsible for driving down nurses' wages or creating shortages. Two methods have been used to identify monopsony power in a labor market. The first is to exploit cross-sectional variation in hospital concentration and wages across local labor markets. Following this method, Hirsch and Schumacher (1995) use wage data for 252 labor markets from the 1985-1993 CPS surveys, along with data on hospitals and market size, to study the relationship between market concentration and wages. They do not find any evidence of a monopsony effect on wages. The second method is to estimate labor supply elasticity at the individual hospital level. Sullivan (1989) takes this approach, using data from the 1979 to 1985 American Hospital Surveys, and finds evidence of substantial monopsony power on the part of hospitals. Staiger, Spetz and Phibbs (1999) also find evidence of monopsony power, using a unique strategy to identify wages in the labor supply equation. Their study is a natural experiment, based upon a federally mandated wage increase at approximately two-thirds of all VA hospitals nationwide that took place in 1991. They find that labor supply of nurses at the individual hospital level is relatively inelastic, supporting the idea that hospitals do have market power when hiring R.N.s.[4]

Another possible explanation that has been advanced to explain the low wages of nurses and other types of care workers is related to the caring nature of their work. If workers derive personal, emotional or spiritual satisfaction from caring for others, then a model of compensating wage differentials might predict that these workers' reservation wages would be lower for care jobs than for other jobs requiring similar skills.[5] Folbre and Nelson (2000) note that in the sociological and feminist economics literatures, the idea of providing care to the elderly and children through a market mechanism has been regarded with some suspicion. The concern is that paying for a "commodity" that is so inherently personal both devalues it from society's perspective and lowers its quality. Interestingly, Heyes (2007) presents a very similar argument in the context of a theoretical economic

[4] Monopsony might also be an explanation for direct care work wages and shortages, although the set of potential employers for these workers is slightly more diverse. However, it will not be possible to test the monopsony hypothesis in this paper because the primary data set (SIPP) does not have specific enough geographical identifying information.

[5] One could also argue that there would be a positive compensating differential for these jobs because of the extremely high risk of on-the-job injury. The non-fatal injury rate in U.S. nursing homes in 2005 was 9.1 per 1,000, well above the private sector average of 4.65 (BLS 2005). See Myers et al. (1993) and Personick (1990) for detailed discussions of on-the-job injuries in this occupation group.

model. In his model, attracting nurses to the profession with higher wages lowers average quality of nursing care by promoting the selection of workers into the occupation who do not the have personal motivations for caring that make a “good” nurse. However interesting, hypotheses about caring externalities are extremely difficult, if not impossible, to test in wage or employment data.

Finally, several studies have considered the role of property rights in determining the wages of nursing home workers. Borjas et al. (1983) find that workers in for-profit nursing homes earn lower wages than workers in non-profit homes, even after controlling for worker characteristics, using the 1973-74 National Nursing Home Survey. Further, the flat rate reimbursement scheme of the Medicaid program put a stronger downward pressure on wages in for-profit homes, compared to non-profit and government homes. Preston (1988) finds a similar non-profit effect in day care centers, but only in the federally regulated sector of the market. Leete (2001) examines non-profit wage differentials in a variety of different industries, using cross-sectional data from the 1990 PUMS. Although the aggregate wage differential is approximately zero, workers in care giving sectors (including nursing and child care) do have significantly higher wages in non-profit establishments. According to Leete’s Census data, more than 80 percent of nursing and personal care jobs were with for-profit firms, so the interaction of competitive pressure on wages with low Medicaid reimbursement levels (which will be discussed in Section 4) may be an important piece of the puzzle in explaining direct care wage levels.

B. Labor supply

Most of the attention devoted to nursing labor supply in the economics literature has focused on the market for R.N.s. Some of this research has been done in the United States, but much has also been done in the United Kingdom and Scandinavian countries, where nursing shortages have been more widespread in recent decades. Antonazzo et al (2002) provide an excellent summary of what has been learned in the nursing labor supply literature. The authors report on a number of studies of labor force participation and hours decisions. Several themes emerge from this summary. First, in many studies, presence and age of children and other demographic characteristics of a nurse and her family played as important a role in employment decisions as economic variables. Second, methodological improvements over time have yielded less elastic estimates of labor supply. Labor supply elasticities in excess of one were not uncommon in early studies, but there was nothing approaching a single consensus estimate (although own-wage elasticities were almost exclusively positive). However, in the past decade, as more studies have controlled for wage endogeneity, unobserved heterogeneity and accounted for full-time versus part-time employment, elasticities have fallen dramatically. A study by Askildsen et al. (2003), using a panel of Norwegian nurses and finding an own wage elasticity of 0.21, is representative of this more state-of-the art work.

Parker and Rickman (1995) provide a model of employment transitions for R.N.s. Using linked data from the 1980-1990 Current Population Surveys to estimate one-year occupational exit rates, they find that while the rate at which R.N.s switch to other occupations remains constant over the period, the rate at which they withdraw from the labor force increases from 13 percent to 18.5 percent per year. As expected, wage is a significant predictor of remaining in the occupation, with an elasticity of 2.00 for married nurses, although the magnitude of the effect ($\eta = 0.13$) is much smaller for single women. The only potential drawback to using the CPS for an employment duration analysis is that the data only capture changes between years and do not allow for the measurement or analysis of shorter spells.

The only paper that, to our knowledge, separately identifies nursing aides in a labor supply analysis is Burkett (2005), who estimates time series models of wages and employment, using aggregate data on nursing aides and R.N.s for 1987 to 2002. Wage elasticities of labor supply are positive and significant for both groups in the baseline estimates, but much larger for nursing aides (2.3) than for R.N.s (0.6). The Bayesian elasticity estimate for nursing aides is lower, but still surprisingly large at 1.6.

4. Medicaid and long-term care

Public policy, particularly the Medicaid program, plays a major role in the market for long-term care in the United States and, consequently, has the potential to affect the direct care labor market. Although the Medicare program only covers shorter stays in skilled nursing or rehabilitative facilities, Medicaid covers nursing home expenses for the low-income elderly and an increasing number of middle-income Americans who "spend down" assets in order to become Medicaid-eligible. As Figure 4 shows, Medicaid was responsible for 53 percent of nursing home expenditures in 2004. Although the breakdown of expenditures between out-of-pocket, private insurance and Medicaid has remained roughly constant since 1990, the absolute level of these expenses has doubled. Medicaid payments to nursing homes have increased from $24 million to $52 million in just 14 years.

Medicaid covers nursing home stays in all states, some type of home health service in most states and the services of personal or home care aides in approximately half of the states (Kaiser Family Commission 2007). Long-term care spending therefore represents a significant fraction of total Medicaid spending at both the federal and state levels. Although the elderly represent only 9.5 percent of Medicaid enrollees, expenditures for this age group make up more than 26 percent of Medicaid spending (Green Book 2004). Appendix Table 1 contains total Medicaid expenditures on both nursing homes and home health services per over-65 resident for 2003 as well as a list of states that covered personal or home care services in 2002. Average annual expenditures per over-65 resident on nursing homes rose from $1,046 in 1996 to $1,244 in 2003. Average spending per over-65 resident on home health rose even more dramatically over the same period, from $98 to $232. There is significant variation in spending patterns across states, with Medicaid nursing home spending ranging in 2003 from less than $500 per year per over-65 resident in Oregon, Nevada and Arizona[6] to more than $3,700 in New York. Similarly, 21 states spend less than $100 per year per over-65 resident on Medicaid home health, while nine states spend more than $400.

One of the effects of Medicaid's role as a primary funding mechanism of long-term care is that low and fixed reimbursements for the majority of patients constrains the ability of nursing homes and home health agencies to offer an adequate wage to attract workers. For this reason, 23 states have adopted wage "pass-throughs" for direct care workers into their Medicaid policies. These programs are a new phenomenon; all but one of the states implemented their pass-throughs after 1998. Appendix Table 1 provides a list of states that had pass-through provisions by the end of our sample period. It is worth noting that this list includes a geographically and politically diverse set of states and that implementation of a pass-through does not appear to be strongly correlated to generosity of other health or

[6] Arizona is unique in that it offers all Medicaid long-term care through a comprehensive Home and Community Based Services (HCBS) 1115 waiver.

welfare policies. States implementing this policy are not just the traditionally more progressive states, such as Massachusetts and California, but also Texas and South Carolina, which are traditionally more conservative in social spending. As Fitzgerald (2001) notes, this may be due to the fact that the political constituency for wage pass-throughs, which includes middle and high-income seniors and their children, is much different than that for other policies affecting low-wage workers, such as the minimum wage or TANF.

The first pass-through provision was passed in Michigan in 1989 and was put into widespread effect during fiscal year 1992. The structure of the Michigan pass-through is one of the most common across all states with these policies. The law states that increased funding up to the equivalent of 50 cents per employee hour will be made available to qualified, Medicaid-reimbursed nursing homes and that these funds are to be used to increase wages and benefits including health insurance, retirement and child care. Nursing homes are required to document the use of the funds. In 2000, the pass-through money was to come out of a 5.4 percent overall increase in state Medicaid funding for long-term care services. (SEIUMI 2006). Because Michigan had passed a pass-through provision in all budget years between 1992 and 2000, the effect on direct-care compensation was cumulative, totaling $4.12 per hour by 2000 and translating into an increase in hourly wages of nursing home aides of $2.05 per hour between 1995 and 2000. (Walker et al. 2000).

Other states set up their pass-throughs in slightly different ways. Minnesota, for example, increased its nursing home reimbursement rate by 40 percent in 1999 and mandated that 80 percent of the increase be used by the homes for wages and benefits. (PHI 2003). Similarly, California passed a $35 million increase in funding to long-term care in the Medi-Cal program in 1999 ($130 million in 2000), to be distributed to all certified nursing homes in the state for the purpose of increasing employee compensation. (Carlson 2002). In both of these states, all nursing homes were expected to participate in the program, but incentives for compliance were weak and participation in the program was likely to have been affected by state auditing practices. Many policy analysts, including Carlson 2002 and PHI 2003, have suggested that auditing is a significant determinant of the effectiveness of the provisions. Although this is an important concern, assessing the implementation of the policies on a national level requires data that is not yet available and is beyond the scope of this paper. However, it remains a goal for future work in this area.

Figures 5 and 6 show trends in Medicaid nursing home and home health spending, respectively, from 1996 to 2003. Spending trends are presented separately for states with and without wage pass-throughs in effect by 2002.[7] In Figure 5, there is modest growth in nursing home spending overall; although the level of spending is consistently lower in states that adopt pass-through programs, there is no significant difference in trend between pass-through and non-pass-through states over the period. However, patterns for home health spending in Figure 6 tell a different story. There is more growth overall in home health spending and the rate of growth appears to increase in 2000. Further, although spending levels are lower in pass-through states in 1996, the rate of spending is increasing relative to non-pass-through states, particularly after 2000, so that pass-through states spend slightly more on home health per over-65 resident by 2003. We will take into account this observable correlation between pass-through status and home health spending in the empirical analysis presented in Section 6.

5. Data

The individual-level data used in this analysis come from the 1996 and 2001 Surveys of Income and Program Participation (SIPP), panel studies that follow separate, nationally representative samples of Americans from 1996-2000 and 2001-2003, respectively. Respondents in each SIPP panel are surveyed every four months and are asked at each interview to provide month-level employment, family structure and other information about the most recent month, as well as about the three preceding months. These four-month reference periods are called waves. For individuals who participate in the survey the entire period, this results in a 48-month panel of continuous demographic and employment data in 1996 and a 36-month panel in 2001. Although these relatively short panels have the disadvantage[8] of not completely capturing employment spells of more than three or four years, by providing monthly data they allow us to measure shorter job spells for lower-skilled women with a much greater degree of accuracy. An additional advantage of the SIPP

[7] With the exception of Michigan, all states that implemented pass-throughs during this period did so between late 1999 and 2002.
[8] This would be relative to an annual observation panel data set, such as the Panel Study of Income Dynamics (PSID) or the National Longitudinal Survey of Youth (NLSY).

data is that respondents are asked a complete set of questions each month on up to two different jobs held that month.

The SIPP codes both industry and occupation group at the three-digit level. We use the occupation code to identify direct care workers as those belonging to the "nursing aide or orderly" and "health aide, not nursing" groups. We further restrict the sample of direct care workers to those in the health care sector and to women.[9] We identify 3,613 spells of direct care work for 2,875 different women by this method. Of the 618 women with multiple job spells, 412 have two spells, 121 have three spells and only 85 have more than three spells.

Table 1 provides descriptive characteristics for these spells, with all monthly observations summarized in the first column and the first month of a given direct care job in the second. These figures show that the direct care workforce is approximately one-third black and more than 10 percent Hispanic, and that direct care workers are less likely than the female workforce in general to be married but more likely to have children. Perhaps unexpectedly for a group of workers with average monthly earnings of only $1,433 overall, educational background is quite varied. While 18 percent of workers starting direct care jobs do not have a high school diploma, 45 percent have at least some education beyond high school and 17 percent have either an associate's or bachelor's degree. Consistent with the distributions in Figure 1, the workers are employed in a variety of industries, with 22 percent working in hospitals, 34 percent working in nursing homes, 36 percent working in other health settings and 8 percent working in community care.

Wages are imputed for each month using monthly earnings, weekly hours and weeks worked per month that are specific to that job.[10] Figure 6 shows the distribution of wages. Table 2 presents descriptive statistics for wages, again separated by all monthly observations and for only the first months at a given job. Not surprisingly, wages at the start of a job spell are approximately $1.00 per hour lower at all points in the distribution. Median (mean) starting wage in a direct care job is $7.47 ($9.63). Approximately one-quarter of the imputed wages fall below the federal minimum wage for most of this period, which is

[9] More than 90 percent of direct care workers are women and we believe that most of the men in these occupations are hospital or residential home orderlies, who usually have a different set of responsibilities than direct care workers, so we restrict our sample to women.

[10] We drop 96 spells that contain months with imputed wages over $50/hour. This exclusion does not change the results presented in Sections 6-8.

possible because the Supreme Court recently upheld a ruling that denies home health and personal care aides working in private households overtime and minimum wage coverage under the Fair Labor Standards Act (Greenhouse 2007). This table also presents summary statistics for the difference in hourly wage between the first and last month spent at a given job. Wage growth appears to be small; the median difference is equal to zero.

We construct measures of the length of job spells by making use of the fact that the SIPP asks a comprehensive set of employment questions at every wave of data collection and for every month within a wave. Employed individuals are given identification numbers that are unique to each employer they have; changes in this number from month to month allow us to track job changes. Continuity from month to month in the employer ID is our measure of job spells.[11] However, because respondents are asked about monthly changes in employment status but only interviewed every four months, a potential weakness of the data is "seam bias" in reporting. And, in fact, we do observe a bunching of employment transitions in the interview months. Following Shore-Sheppard et al (2007), we use all monthly observations in our analyses of employment duration and control for seam bias with a simple "seam month" dummy variable.

Table 3 contains detailed descriptive statistics for the spells of direct care employment and Figure 7 depicts the distribution of these spell lengths in our sample. We break the spells down by censoring status. Approximately one-third of spells are non-censored, meaning that they started after the SIPP panel began and ended before the panel ended or the respondent attrited from the survey. Of the remaining spells, 719 are left-censored (reported in the first month of the panel and presumably ongoing before the survey), 919 are right-censored (reported in the last month of the panel and presumably ongoing after the survey) and 890 are both right and left-censored. As would be expected, the non-censored spells are much shorter, with mean (median) length of 5.36 (four) months compared to 15.76 (10) months for the right and left-censored spells.[12]

[11] For each job held, a start date is also collected and an end date is recorded in the month that the individual reports leaving the job. However, because this measure includes spells that began before the SIPP panel started, we are not able to control for family structure, asset income and other factors when the spell started, and so we use employer-ID based job spells as the measure of employment duration.

[12] One of the complicating factors in the SIPP data is that a small number of respondents have missing months (or, more commonly, an entire wave). We drop all observations for individuals who appear to leave and then return to the panel so that we do not mis-measure spells in the missing period.

Finally, during the first interview of each panel, the SIPP asks respondents who are working how many total lifetime months they have worked in their current primary occupation. While this one-time question does not provide enough variation for a multivariate analysis of occupational tenure, it does provide some information about direct care workers' total experience in these jobs. We have 532 observations for individuals working in direct care as a primary job in March 1996 and 669 observations in January 2001. The average length of time in the direct care occupation is approximately nine years in 1996 and 10 years in 2001. Median lengths are somewhat shorter. Since the numbers themselves are not particularly informative, a simple comparison may be helpful. The mean length of time in an occupation for a direct care worker is significantly shorter than that reported by R.N.s and longer than that reported by child care workers, but the estimates are almost identical after controlling only for the age of the workers. This may not be surprising given the similarities between these occupation groups, with labor shortages in the market for R.N.s and much lower wages and turnover in the market for child care workers.[13]

6. Wages

We first turn to an analysis of the determinants of hourly wages for direct care workers. Wage models are estimated on both a sample of all monthly observations and on a sample of just the first month of each spell by OLS. The wage equations take the following form:

$$Ln(wage)_{jt} = \alpha_s + \alpha_y + X_{jt}\beta + Z_{st}\gamma + J_{jt}\psi + \varepsilon_{jst} \qquad (1)$$

where the j subscript refers to a job spell, since some individuals have more than one job spell in the panel. The vector X includes a full set of demographic characteristics, including age and its quadratic, race/ethnicity, metropolitan area residence and education. This model also includes a set of indicators for job characteristics in the vector J, including whether the worker is covered by a union contract, whether the employer has more than 100 total employees and the sector of employment (hospital, nursing home, and community care with other health as the omitted category). The vector Z contains state policies and economic characteristics, including Medicaid spending on nursing homes and home health services

[13] Child care workers are the only occupational group that we have identified in the SIPP data that have significantly shorter employment spells than direct care workers. These results are available upon request.

per over-65 resident, the unemployment rate, minimum wage, fraction of population in the over-65 population groups, the federal Medicaid matching rate (FMAP) and a dummy variable for a state having a wage pass-through program in effect in a given year.

One issue in estimating these models is that the number of observations may overstate the true amount of variation in the data. In these data, this is a particular concern because of the expected impact of state-level policy variables. Additionally, error terms are likely to be correlated across individual or job spell in the all-months models or cases of individuals with more than one spell in the data panel. In these data, all multiple layers of clustering (individual, job, state) are nested in that none of the direct care workers in the sample move between states. Because Cameron et al (2006) and Pepper (2002) show that in this case, it is appropriate to cluster at the highest level of aggregation, we cluster standard errors in all of our models by state.

The first two columns of Table 5 present the results for the wage models. The sample in Column 1 contains all monthly observations and the sample in the Column 2 contains only the observation for the first month of a job spell. For many of the variables, the results in Columns 1 and 2 are qualitatively similar, although fewer results are significantly significant in the smaller-sample first month equation. Below, we discuss the results for the full sample (Column 1).

A number of worker demographic characteristics are significantly associated with hourly wage. There is a positive and diminishing effect of age on wage, averaging approximately 2 percent per year. Family status is also significantly correlated with wage. Although children do not have a significant effect on wage, married women earn 4.7 percent per hour less than single women. This may reflect married women's ability to be more selective in the wage offers they accept. Ethnicity also appears to be very strongly correlated with wage. Although there is no statistically significant difference between the wages of non-Hispanic whites and blacks, Hispanics earn 8.3 percent less per hour than non-Hispanic whites. The SIPP does not contain citizenship information in the monthly (core) file, and so this coefficient is likely to be picking up the lower wages earned by non-native workers, particularly in home health settings. Where a direct care worker lives also significantly affects wages; workers living inside of Metropolitan Statistical Areas (MSAs) earn 7.5 percent more than their rural counterparts.

Increasing education at every level also has large and statistically significant effects on wages. Non-high school graduates earn 8.6 percent less per hour than those with high school diplomas. Education beyond high school also significantly increases wages, with the effects ranging from 4.1 percent for workers with some college but no degree to almost 30 percent for workers with bachelor's degrees. These strong effects of education on wages are consistent with the predictions of a general human capital model, but are perhaps surprising within such a narrowly defined occupational group, especially when sector of employment is included in the model. Our tests (results available upon request) indicate that the effect of education on wage is not confined to any single employment sector. The most likely explanation is that there are unobservable differences in skill requirements even within occupation and employment sector.

Perhaps surprisingly, there is no statistically significant effect of either union coverage or firm size on wages of direct care workers. However, sector of employment is strongly correlated with hourly earnings. Although the wages of workers in nursing homes, home health (the omitted category in these models) and community care are not statistically different, nursing aides in hospitals earn a full 25.3 percent more per hour than other direct care workers, even after controlling for narrow education categories. One possible explanation is that the hospital coefficient is picking up the effect of the dramatically larger size of hospitals as employers. The large hospital wage premium may also reflect either a higher level of (unobservable) human capital required of hospital workers or a compensating differential for unattractive features of hospital jobs.[14]

Of the economic and policy variables included in the model, only three have statistically significant effects. As predicted, Medicaid wage pass-throughs are associated with a significant 7 percent increase in wages for direct care workers. The effect of the minimum wage on direct care worker salaries is unexpectedly negative, indicating that direct care workers in states with high minimum wages actually earn lower wages than direct care workers in other states. The most logical explanation for this result would be that a state's minimum wage changes the pattern of selection into direct care jobs. We further

[14] Although nursing aides in and out of hospitals perform most of the same tasks, the patients in a hospital setting are more seriously ill and may require more careful attention. And despite the fact that hospital and other direct care workers are about equally likely to work full-time schedules (Smith and Baughman 2007), hospital workers may receive a compensating wage differential for the less flexible scheduling that hospital work offers.

discuss this possibility below. The fraction of a state's population over age 65 also has an unexpected negative effect on wages, with an elasticity of close to -1.0. All else being equal, we would expect increases in this variable, as a proxy for demand for long-term care services, to result in higher wages. However, it may be an imperfect proxy for demand.

Overall, we find that many of the demographic factors one would expect to influence wages, including age, area of residence, ethnicity and education, do have significant effects. Additionally, despite the relatively similar job descriptions of direct care workers, sector of employment has a large effect on hourly earnings, with hospital aides earning much more than workers in other sectors. Finally, the only public policy that has been created specifically to address low wages in the direct care workforce, the Medicaid wage pass-through, does appear to be working. Direct care workers in states with pass-through programs earn 7 percent more per hour than workers in other states.

There are also several significant results of the full-sample wage model in Column (1) that, theoretically, would not necessarily be expected, including the positive effect of marriage and the negative effect of state minimum wage. Both of these results suggest that selection into direct care occupations might be non-random with respect to the covariates in the model. We have explored this possibility by estimating several specifications of occupational selection models and a two-stage Heckman wage model with selection, but we do not find any evidence of selection patterns that would bias our results.

7. Health insurance coverage

Another source of compensation for direct care workers is health insurance coverage. Only 57.3 percent of direct care workers in the SIPP sample are covered by employer-sponsored health insurance in any given month, compared to 69.4 percent of female workers with less than a college education and 65.1 percent of female workers earning less than $12 per hour. And health insurance may be particularly important in this workforce, given both low average wages and high occupational injury rates (Personick 1990 and Myers et. al 1993).

We estimate a linear probability model of employer-sponsored health insurance (ESI) coverage our sample of direct care workers, using a model almost identical to our wage model.

$$ESI_{jt} = \alpha_s + \alpha_y + X_{jt}\beta + Z_{st}\gamma + J_{jt}\psi + M_{st}\delta + \varepsilon_{jst} \quad (2)$$

The dependent variable is a discrete measure of health insurance coverage through a current employer, which in this case can be the employer in either a primary or secondary job. The set of independent variables is identical to the one in the wage equation, with the addition of the vector M, which contains measures of two state-level policies that have been shown in previous research to affect employer-sponsored health insurance coverage rates.

The first variable in M is a simulated instrument for child Medicaid eligibility, which is constructed by taking a nationally representative sample of children under age 19 from the 1996 SIPP and using eligibility data from the National Governors' Association to calculate the fraction that would be eligible for Medicaid or State Child Health Insurance Programs (SCHIPs) under each state's rules for a given year. This measure varies within and across states over time, based upon the legislative generosity of a state's Medicaid policy toward low-income children.

Because of the correlation between eligibility for infants and pregnant women in Medicaid and children and their parents in some state SCHIP programs, this measure should also be a good proxy for coverage of low-income women. The other variable in M is also a simulated instrument for state personal income taxation that is calculated using the TAXSIM model (Feenberg and Coutts 1993). Employer-sponsored health insurance is not taxable for purposes of federal or state income taxes, and a number of studies have found a significant relationship between marginal income tax rates and rates of employer-sponsored coverage (see, for example, Gruber and Poterba 1994 and Royalty 2000). Our tax measure is a state's average marginal income tax rate calculated on a 1995 national sample of households using that state's tax rules for each year.[15]

We estimate this model of health insurance coverage on the full sample of all monthly observations (Column 3) and on the sub-sample of first months of each job spell (Column 4). The results are generally consistent in sign and significance, so we will focus the discussion below on the full-sample results in Column 3.

Many of the demographic variables that affected direct care worker wages also affect health insurance coverage. Older workers are significantly more likely to receive health

[15] See http://www.nber.org/~taxsim/state-marginal/ for a full description of this variable.

insurance from their employers, as are married workers. However, conditional upon marital status, workers with children are less likely to be covered by ESI. This may reflect different rates of participation in employer-sponsored programs by mothers compared to childless women, rather than differences in the rates at which the two groups are offered insurance by their employers. Note that our dependent variable measures whether a worker actually participates in an ESI plan; some workers who do not have ESI may actually be offered it (at some price) and decline. Many direct care workers have earnings and family incomes low enough to qualify their children for Medicaid or SCHIP health insurance coverage and this may lead them to decline ESI altogether.[16]

Educational attainment is also significantly related to health insurance coverage, although not as strongly as it is to hourly wages. The largest effects are at the lowest education levels. Not having a high school degree is associated with a significant 9.6 percentage point (or 18 percent) lower likelihood of health insurance coverage. There is also a marginally significant 6.4 percentage point (10 percent) increase in coverage for workers with a trade degree or diploma compared to workers with only a high school degree.

As was true with wages, place of employment also has significant effects on the probability of employer-sponsored health insurance coverage. Consistent with standard predictions about risk pooling in insurance markets, we see that those who work for large employers are 5.7 percentage points (9 percent) more likely to have ESI and this is true even while controlling for sector of employment. There are also independent positive effects of employment in hospitals and nursing homes. Workers in hospitals have 31 percentage points (43 percent) and workers in nursing homes 11.3 (18 percent) more likely to have ESI coverage than workers in the reference home health sector.

Finally, few of the economic and policy variables in the model have statistically significant effects on employer-sponsored health insurance coverage. Workers in states with higher fractions of employment in the manufacturing industry are more likely to have coverage. This may reflect a spillover effect of high coverage rates in manufacturing jobs (and more influence of labor unions) in areas where this is the predominant source of lower-skill employment. The only other significant variable is a large negative effect of the fraction of population over 65 on ESI coverage. As in the wage model, this result is puzzling.

[16] Some states, such as New York, also offer health insurance to the parents of eligible children under their SCHIP programs.

Surprisingly, neither Medicaid eligibility nor state personal income tax rates are significantly associated with rates of employer-sponsored health insurance coverage in these models.

8. Employment duration

The final step of this labor supply analysis is the estimation of employment duration models. We estimate the probability of leaving a job with a given employer as a discrete hazard using a stacked logit model. The logit model takes the following form:

$$P(END)_{jt} = \alpha_s + \alpha_y + \beta_1 \ln(w_{jt}) + \beta_2 ESI_{jt} + X_{jt}\delta + J_j\psi + Z_{st}\phi + \gamma_1 Seam_{jt} + \gamma_2 PD_{jt} + \varepsilon_{jst} \quad (3)$$

where for each job spell j the subscript t varies from 1…τ where τ is the last observed month in a spell (which may be either a transition month within the panel or a censored observation at the end of the panel). The dependent variable (END) is coded as 1 only for observed transitions; right-censored spells are treated as ongoing. The primary independent variables are log wage and ESI coverage. The X vector contains measures of age, child age, marital status, other family income, race, urban/rural residence and education and the J vector contains the same union, employer size and employer type variables in Equations 1 and 2. The vector Z contains three state-level variables that capture some of the opportunities available outside of a direct care job: the minimum wage, unemployment rate, and AFDC/TANF benefit available to a family of three. Seam is a dummy variable that marks observations for interview months (occurring every fourth months) and controls for "seam bias" effect of a disproportionate number of transitions reported in these months. The period (PD) variables are dummy variables for each wave (1-12 in the 1996 panel and 1-9 in the 2001 panel). Therefore, the γ_2 coefficients in this model non-parametrically estimate a baseline hazard that is unrelated to the effects of other covariates in the model, controlling for the possibility of duration dependence in job spells.

The results of the baseline logit models are presented as odds ratios in Columns 1 and 2 of Table 6. Values above one indicate that a variable increases the probability of leaving a job in a given period and values below one indicate that a variable decreases this probability. Because spells that are ongoing when the survey starts (left-censored) are expected to be inherently different than new spells (and because there is no accepted technique in hazard modeling for dealing with left-censoring), we separately estimate models for spells with and without left censoring.

As expected, higher wages significantly decrease the probability of a direct care employment spell ending in any given month. The odds ratio associated with a 100 percent increase in wage is 0.78 in Column 1 and 0.76 in Column 2, suggesting duration elasticities of approximately 0.20. Health insurance coverage provided by an employer also significantly lengthens job spells, with odds ratios of 0.58 in Column 1 and 0.52 in Column 2. However, the wage effects are likely to be biased if wages increase as job spells go on and workers acquire more tenure and experience. Similarly, some employers require a waiting period before workers qualify for ESI plans, so in the case of health insurance, there is also likely to be a spurious correlation between coverage and employment duration. To get around this endogeneity problem, we re-estimate the logit employment duration model, substituting wage and ESI status in the *first* month of each job spell for the value in all other monthly observations.[17] The wage effects in these models, reported in Columns 3 and 4, are still significant although, as expected, slightly smaller. Odds ratios of 0.90 and 0.85 suggest duration elasticities of approximately 0.15. The magnitude of the health insurance effect also shrinks, although the effects are still quite strong with odds ratios of 0.62 and 0.70 in Columns 3 and 4, respectively. This translates into direct care workers being 30 to 38 percent less likely to leave a job in a given month if that job provides their health insurance coverage.

There are a number of other variables in the model that are significantly associated with length of job spell, even after controlling for wage and health insurance coverage. Younger direct care workers are more likely to leave a job in a given month. Workers with children under age 5 are 26 percent more likely to experience job transitions compared to workers with children older than 10, although this effect is only significant in the non-left censored sample. Workers in non-rural areas are 25 percent more likely to leave a job in a given month, but this effect is only significant for left-censored samples. The main effect of education is a 36 to 46 percent higher probability of leaving a job in a given month for individuals with trade school degrees; the effect is stronger for left-censored spells. There is

[17] Our preferred method for dealing with endogeneity would be to instrument for each month's wage and ESI status. However, we have a weak instruments problem in that the R^2 in our instrumenting equations is low in all specifications and of the policy variables, only wage pass through is significant and of the expected sign.

also a 19 percent higher monthly transition probability for workers with some college relative to a high school diploma. These effects may be explained by some workers acquiring the Certified Nursing Aide (C.N.A.) certification, which opens up a number of job opportunities in nursing homes and hospitals that would not be available to home health or personal care aides.

Place of employment also affects job spell length in certain cases. Even after controlling for wages and health insurance, longer employment duration is associated with larger employers (only significant in the non-left censored sample) and hospital employment (only significant in the left-censored sample). These variables may be capturing other fringe benefits offered by larger employers besides wage and health insurance, including public transit assistance or parking, day care assistance, meals or paid vacation time.

Of the economic and policy variables, only state minimum wage appears to affect exits from direct care jobs and the effect is only significant in the non-censored sample. After controlling for own wage, a higher state minimum wage is associated with longer job spells and the magnitude of the effect is surprisingly large. This, like the effect of minimum wages on hourly wages in Section 6, is still something of a puzzle. Finally, we find that, as expected, exit rates from direct care jobs are significantly higher for secondary jobs and during the seam months of the SIPP.

9. Discussion

The goal of this paper has been to provide an overview of the market for direct care workers in the United States. Given the dramatic increase in population aging that is expected as the baby boom cohort nears retirement age, along with reports of shortage and problematic turnover that already affect the provision of long-term care, this labor market promises to be one of public policy interest in coming years. Using job spell data from the 1996 and 2001 panels of the SIPP, we find both that wages of direct care workers are quite low, with median starting wage of $7.96, and that spells of employment with a given employer are short, averaging just under 10 months.

We find several significant demographic determinants of direct care workers' wages and employer-sponsored insurance coverage rates, including age, family structure, ethnicity, metropolitan area residence and education. Our analysis also reveals significant differences in compensation patterns across sectors of employment, despite the seemingly similar skill requirements of workers in different sectors. Hospitals pay significantly more per hour and are much more likely to offer health insurance to their employees. Nursing home wages are not significantly different from those of in-home health and community care, but nursing home employees are significantly more likely to have ESI coverage. Interestingly, these results hold even when controlling for employer size. Among the policy variables, the main result is that Medicaid wage pass-through programs appear to achieve their goal of increasing direct care worker wages. Direct care workers in states with these programs earn an average of 7 percent more per hour than workers in other states.

This may suggest an avenue for accomplishing a related policy goal: increasing the length of time that direct care workers remain on the job. The effects of starting wage and first-month ESI coverage on the length of the eventual job spell are large and significant for both censored and non-censored spells. While the interpretation of the wage effect on employment duration is straightforward, the interpretation of the health insurance effect is more difficult. In the general labor economics literature, the effect of health insurance in reducing employment transition is termed "job lock" (see, for example, Madrian 1994) and is usually considered to be a source of inefficiency in a dynamic labor market. In the case of direct care workers, there may be a positive efficiency effect of longer job spells if they improve the quality of long-term care.[18] This is an issue that deserves further exploration in future work.

[18] In a model not reported in the paper (though available from the authors by request), we tested the impact of *any* type of health insurance and of public health insurance on job duration, but did not find any significant effects.

References

Ahlburg, D.A. 1996. "The effect of wages on the retention of nurses." *Canadian Journal of Economics* 29: 126-129.

Antonazzo, Emanuela. 2003. "The Labour Market for Nursing: A Review of the Labor Supply Literature." *Health Economics* 12(6): 465-478.

Askildsen, Jan Erik, Badi H. Batalgi and Tor Helge Holmas. 2003. "Wage Policy in the Health Care Sector: A Panel Analysis of Nurses' Labor Supply." *Health Economics* 12: 705-719.

Barry, Theresa et al. 2005. "Nurse Aide Empowerment Strategies and Staff Stability: Effects on Nursing Home Resident Outcomes." *The Gerontologist* 45: 309-317.

Boal, William M. and Michael R. Ramson. 1997. "Monopsony in the Labor Market." *Journal of Economic Literature* 35(1): 86-112.

Borjas, George J., H.E. Frech III and Paul B. Ginsburg. 1983. "Property Rights and Wages: The Case of Nursing Homes." *Journal of Human Resources* 17(2): 231-246.

Burkett, John P. 2005. "The Labor Supply of Nurses and Nursing Assistants in the United States." *Eastern Economic Journal* 31(4): 585-599.

Cameron, A. Colin, Jonah B. Gelbach and Douglas L. Miller. 2006. "Robust Inference with Multi-Way Clustering." NBER Technical Working Paper 327.

Carlson, Eric. 2002. "California's Leaky System to Improve the Wages of Nursing Home Employees." Los Angeles: National Senior Citizens Law Center. March.

Direct Care Alliance. "State Activities to Increase C.N.A. Wages." Accessed 10/26/06 at www.directcarealliance.org/sections/pubs/st_staffing_survey/table1.htm.

Feenberg, Daniel Richard, and Elizabeth Coutts (1993) "An Introduction to the TAXSIM Model." *Journal of Policy Analysis and Management* 12(1): 189-194.

Fitzgerald, Joan. 2001. "Better-Paid Caregivers, Better Care." *American Prospect* online. May 21.

Freeman, Laura. 1995. "Home-sweet-home health care." *Monthly Labor Review* (March): 3-11.

Folbre, Nancy and Julie A. Nelson. 2000. "For Love or Money – or Both?" *Journal of Economic Perspectives* 14(4): 123-140.

Greenhouse, Steven. 2007. "High Court Rules Against Home Aide on Wages." *New York Times* (June 12).

Gruber, Jonathan and James Poterba. 1994. "Tax Incentives and the Decision to Purchase Health Insurance: Evidence from the Self-Employed." *Quarterly Journal of Economics* 109(3): 701-733.

Hernandez-Medina, Esther, Susan Eaton, Donna Hurd and Alan White. 2006. "Training Programs for Certified Nursing Assistants." AARP Working Paper 2006-08. Washington, DC: AARP Public Policy Insitute.

Heyes, Anthony. 2005. "The economics of vocation or 'why is a badly paid nurse a good nurse?'" *Journal of Health Economics*. 24(3): 561-569.

Hirsch, Barry T. and Edward J. Schumacher. 1995. "Monopsony power and relative wages in the market for nurses." *Journal of Health Economics* (October): 443-476.

Hunter, Larry W. 2000. "What Determines Job Quality in Nursing Homes?" *Industrial and Labor Relations Review* 53(3): 463-481.

Kaiser Family Foundation. 2007. State Health Facts Database. Accessed 6/14/07. www.statehealthfacts.org

Madrian, Brigitte C. 1994. "Employment-Based Health Insurance and Job Mobility: Is There Evidence of Job Lock?" *Quarterly Journal of Economics* 109(1): 27-54.

Myers, Ann, Roger C. Jensen, David Nestor and Jacqueline Rattiner. 1993. "Low Back Injuries Among Home Health Aides Compared with Hospital Nursing Aides." *Home Health Care Services Quarterly* 14(2/3): 149-155.

National Governor's Association. Various years. "Maternal and Child Health Updates: State Coverage of Pregnant Women and Children." Washington, DC.

Nelson, Julie A. 1999. "Of Markets and Martyrs: Is it OK to Pay Well for Care?" *Journal of Feminist Economics* 5(3): 1-17.

Paraprofessional Health Institute (PHI) and the North Carolina Department of Health & Human Services Office of Long-term Care. 2002. *Results of the 2002 National Survey of State Initiatives on the Long-Term Care Direct Care Workforce*.

Parker, Carl and Bill Richman. 1995. "Economic Determinants of the Labor Force Withdrawl of Registered Nurses." *Journal of Economics and Finance* 19(1): 17-26.

Pepper, John V. 2002. "Robust inferences from random clustered samples: an application using data from the panel study of income dynamics." *Economics Letters* 75: 341-345.

Personick, Martin E. 1990. "Nursing home aides experience increase in serious injuries." *Monthly Labor Review* (February): 30-36.

Preston, Anne E. 1988. "The Effects of Property Rights on Labor Costs of Nonprofit Firms: An Application to the Day Care Industry." *Journal of Industrial Economics* 36(3): 337-350.

Royalty, Anne Beeson. 2000. "Tax preferences for fringe benefits and workers' eligibility for employer health insurance." *Journal of Public Economics* 75(2): 209-227.

SEIU Michigan State Council. "Wage Pass Through" Accessed 10/26/06 at http://www.seiumi.org/news/wagepass.cfm.

Shore-Sheppard, Lara, John Ham and Xianghong Li. 2006. "Analyzing Movements Over Time in Employment Status and Welfare Participation while Controlling for Seam Bias Using the SIPP." Working Paper, Williams College.

Smith, Kristin and Reagan Baughman. 2007. "Caring for America's aging population: a profile of the direct care workforce." *Monthly Labor Review* 130(9): 20-26.

Staiger, Douglas O., Joanne Spetz and Ciaran Phibbs. 1999. "Is There Monopsony in the Labor Market? Evidence from a Natural Experiment." NBER Working Paper 7258.

Sullivan, Daniel. 1989. "Monopsony Power in the Market for Nurses." *Journal of Law and Economics* 32(2,2): 358-65.

Walker, John S., Steve Angelotti and Dana Patterson. 2000. *Medicaid Funding for Nursing Home Services: A Historical Perspective.* State of Michigan. Senate Fiscal Agency. June.

Wright, Bernadette. 2005. "Direct Care Workers in Long-Term Care." Washington, DC: AARP Public Policy Institute.

U.S. Bureau of Labor Statistics. 2006. "May 2006 National Industry-Specific Occupational Employment and Wage Estimates." http://www.bls.gov/oes/current/oessrci.htm accessed 6/12/07.

U.S. Bureau of Labor Statistics. 2005. "Incidence rates of nonfatal occupational injuries and illnesses by industry and case types, 2005" http://www.bls.gov/iif/oshwc/osh/os/ostb1619.txt accessed 7/30/07.

U.S. Bureau of Labor Statistics. 2004. "2004-14 Employment Projections, Table 3C." http://www.bls.gov/news.release/ecopro.toc.htm accessed 6/12/07.

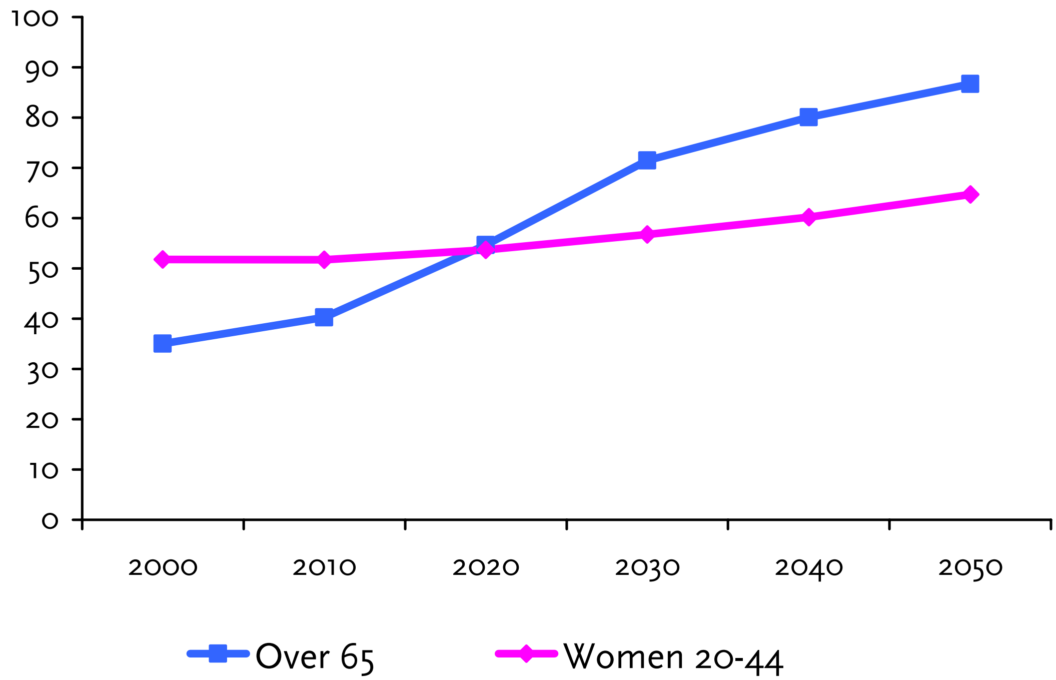
U.S. Congress. Committee on Ways and Means. 2004. *The Green Book*. Government Printing Office: Washington, D.C.

U.S. Department of Health and Human Services. 2002. "State Wage Pass-Through Legislation: An Analysis." (December 20) Accessed at http://aspe.hhs.gov/daltcp/reports/wagepass.htm

U.S. Department of Health and Human Services. Office of the Inspector General. 2002. "Nurse Aide Training." (November) OEI-05-01-00030.

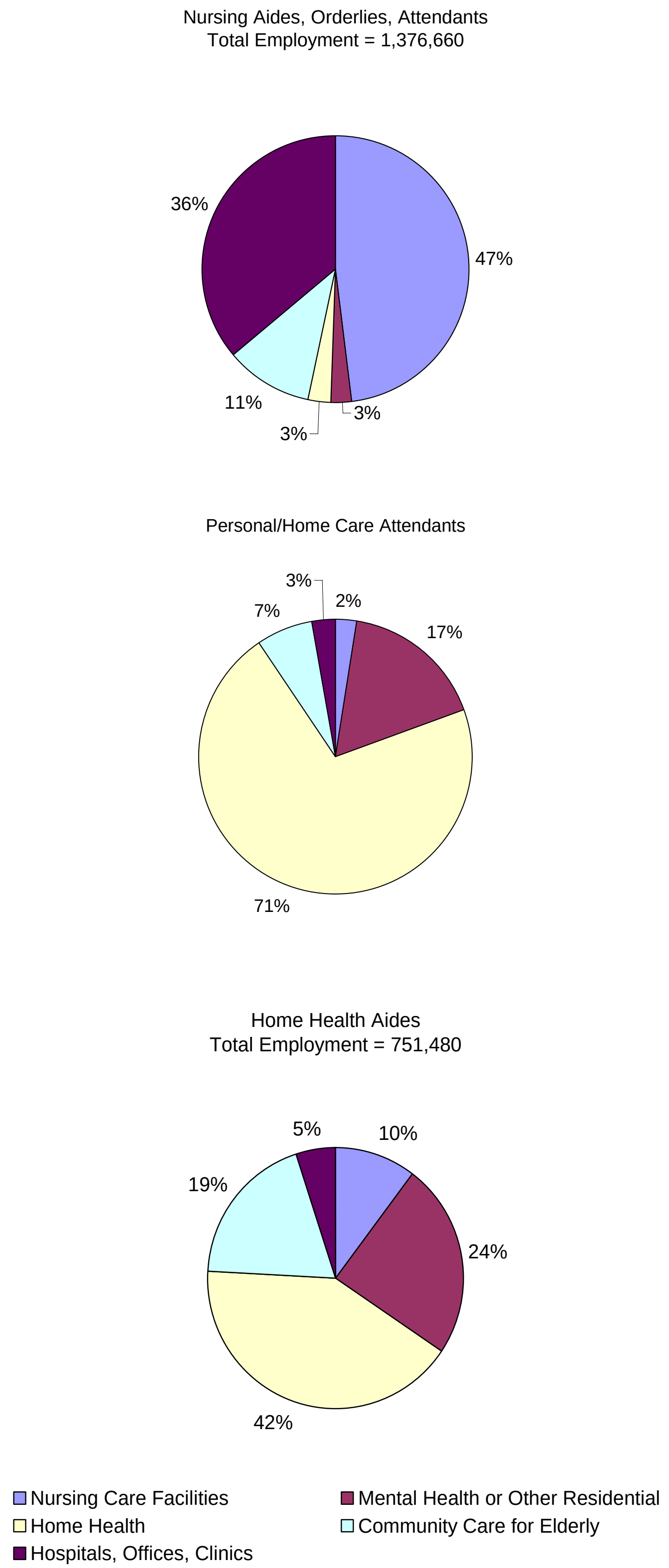
U.S. General Accounting Office. 2001. "Recruitment and Retention of Nurses and Nurse Aides Is a Growing Concern." GAO-01-750T.

Figure 1: Projected Population Growth, 2000-2050



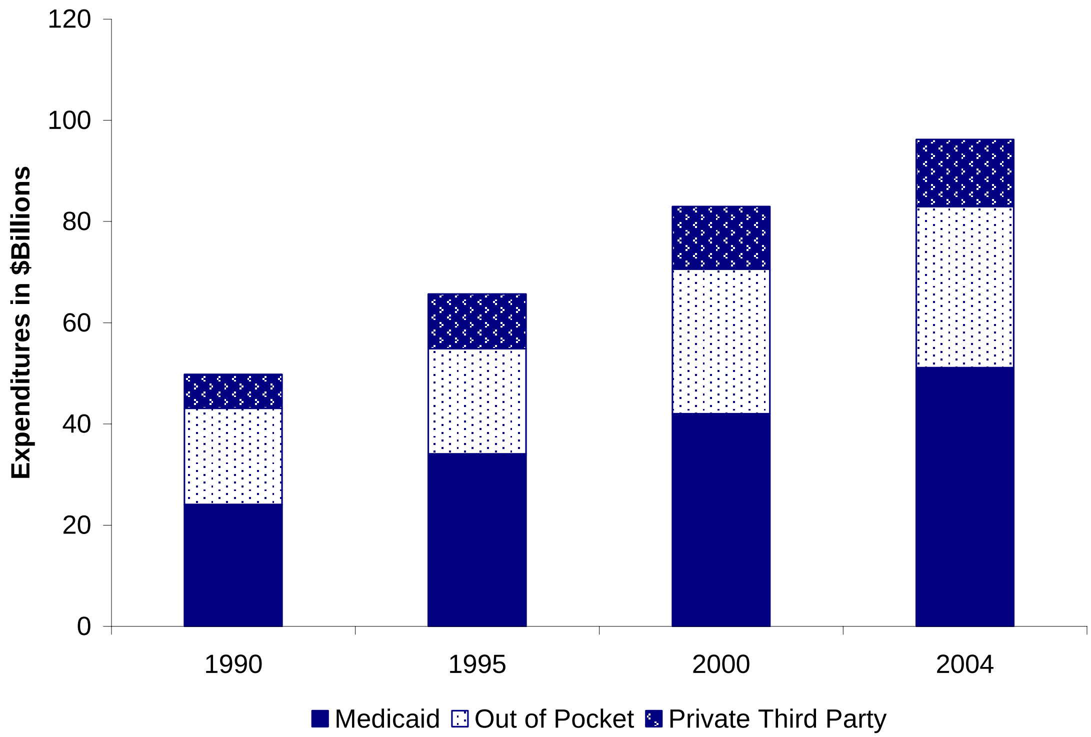
Source: U.S. Census Bureau (2004) population projections

Figure 2: Employment by Occupation and Industry



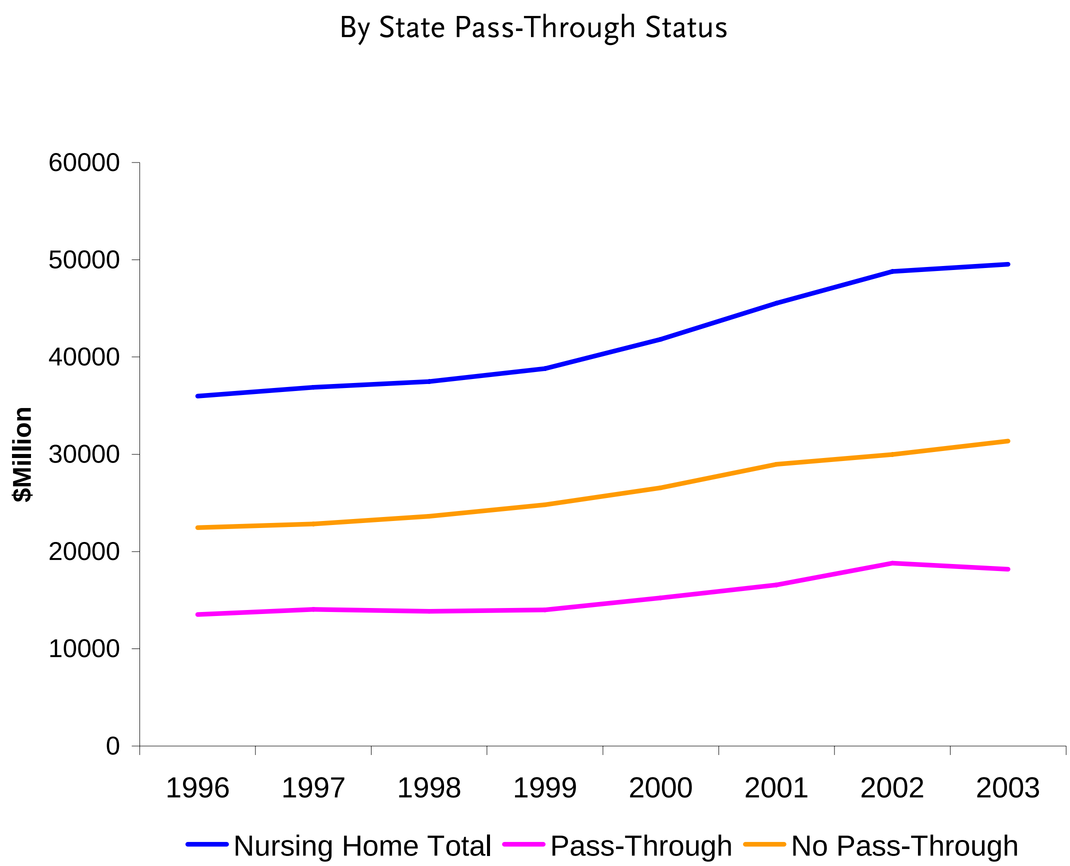
Source: Bureau of Labor Statistics, Occupational Employment Statistics, 2006.

Figure 3: Payments for Nursing Home Care by Source of Payment, 1990-2004



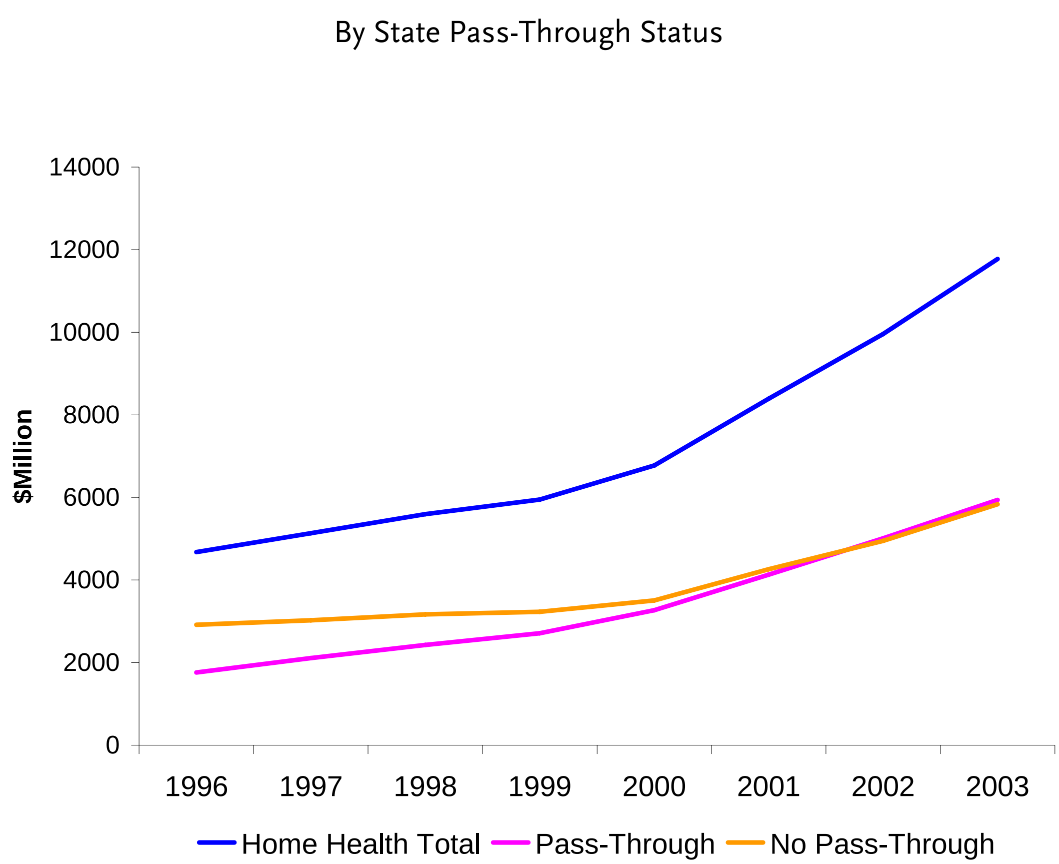
Source: U.S. Centers for Medicaid and Medicare Services. "Hospital and Physician and Clinical Service Expenditures by Source of Payment: 1990 to 2004" [<http://www.cms.hhs.gov/statistics/nhe/default.asp>.](http://www.cms.hhs.gov/statistics/nhe/default.asp)

Figure 4: Medicaid Payments for Nursing Home Care, 1996-2003



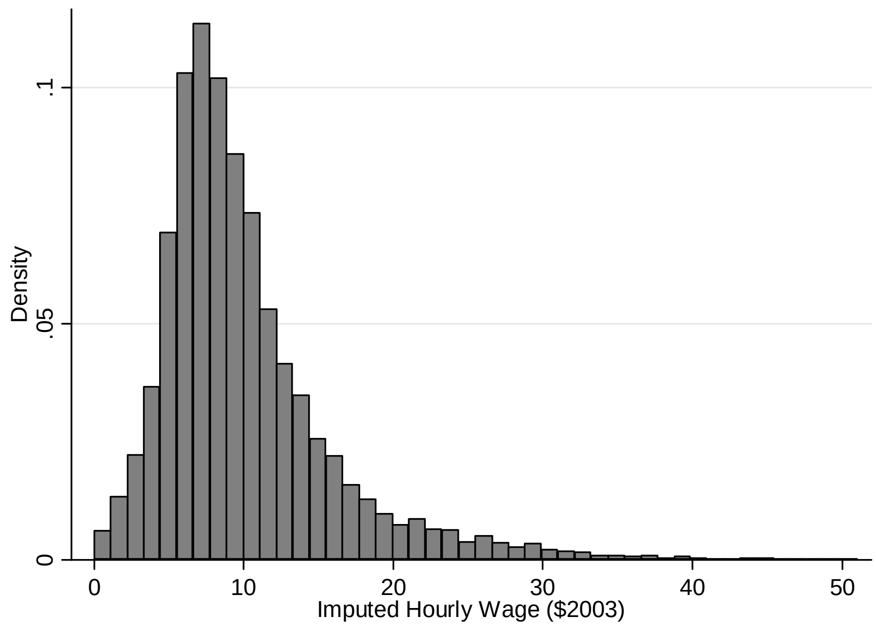
Source: Authors' tabulations from National Health Expenditure Data, Centers for Medicare and Medicaid Services. www.cms.gov

Figure 5: Medicaid Payments for Home Health Care, 1996-2003



Source: Authors' tabulations from National Health Expenditure Data, Centers for Medicare and Medicaid Services. www.cms.gov

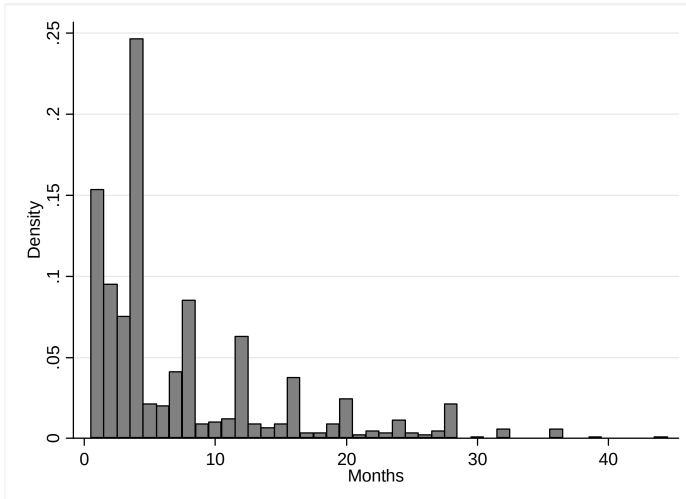
Figure 6: Distribution of Imputed Hourly Wage



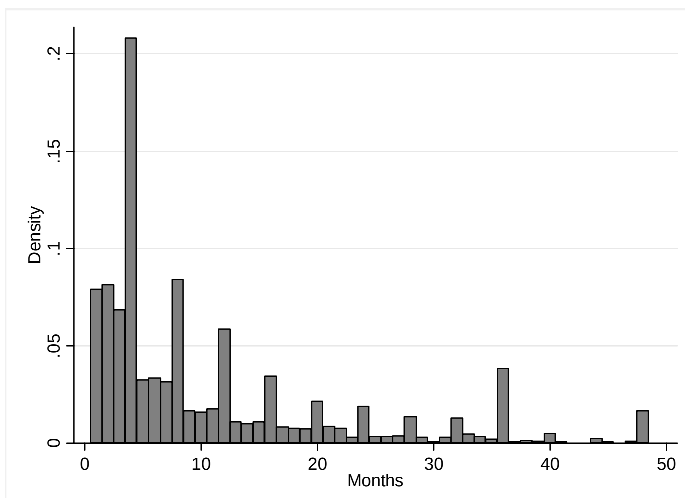
Source: Authors' tabulations on a sample of months of direct care employment (N=x) from the 1996 and 2001 panels of the SIPP.

Figure 7: Distribution of Job Spell Lengths

A. Not Left Censored



B. Left-Censored



Source: Authors' tabulations on a sample of spells of direct care employment (N=x) from the 1996 and 2001 panels of the SIPP.

Table 1: Descriptive Statistics for Direct Care Work Spells

	(1) DCW All monthly observations	(2) DCW Spell first months only	(3) Female workers with < college	(4) Female workers with wage <$12
Individual variables				
Age	42.0 (10.4)	39.8 (10.4)	41.6 (10.0)	41.0 (10.3)
Married	52.6 (49.9)	45.5 (49.4)	60.2 (49.0)	59.1 (49.2)
Hispanic	12.1 (32.6)	11.0 (31.2)	10.3 (30.4)	11.1 (31.4)
Black	31.3 (46.4)	33.1 (47.1)	14.5 (35.2)	14.4 (35.2)
Other non-white	4.3 (20.4)	4.1 (19.9)	3.8 (19.1)	4.4 (20.6)
No children	44.7 (49.7)	40.8 (49.1)	49.7 (50.0)	49.7 (50.0)
Metro area residence	75.3 (43.1)	72.7 (44.5)	61.2 (48.7)	58.2 (49.3)
Less than HS degree	17.1 (37.6)	18.3 (38.7)	11.4 (31.8)	12.8 (33.4)
HS graduate	39.9 (49.0)	36.5 (48.2)	42.1 (49.3)	48.1 (48.5)
Some college	18.3 (38.7)	19.6 (40.0)	26.0 (43.8)	19.9 (40.0)
Trade degree	8.3 (27.6)	8.6 (28.0)	6.9 (25.4)	4.1 (22.0)
Associate's degree	8.8 (28.4)	8.0 (27.1)	13.5 (34.2)	8.4 (27.8)
Bachelor's degree +	7.4 (26.2)	9.1 (28.7)	0 (0)	15.5 (36.2)
Monthly earnings	$1,433 (960)	$1,196 (903)	$1,710 (1301)	$1,146 (631)
Empl. Health Ins. (ESI)	57.3 (49.4)	46.4 (49.9)	69.4 (46.1)	65.1 (47.7)
Monthly other income	$1,840 (2558)	$1,731 (2807)	$2,418 (3054)	$2,430 (3241)
Weekly hours	35.0 (11.3)	33.5 (12.5)	36.3 (10.9)	35.4 (12.0)
Large employer	40.4 (49.1)	38.2 (48.6)	55.5 (49.7)	52.0 (50.0)
Hospital	22.1 (41.5)	17.7 (38.2)	--	--
Nursing home	34.0 (47.3)	36.6 (48.1)	--	--
Other health	35.4 (47.9)	36.0 (48.0)	--	--
Community care	8.4 (27.7)	9.7 (27.6)	--	--
State Variables				
AFDC/TANF benefit	$734 (259)	$747 (259)	$781 (276)	$770 (277)
Minimum wage	$5.27 (0.75)	$5.23 (0.79)	$5.09 (0.77)	$4.11 (0.78)
FMAP	57.70 (7.53)	58.01 (7.39)	57.51 (7.28)	58.02 (7.38)
PC Medicaid NH spend	$1,461 (949)	$1,407 (993)	$1,300 (784)	$1,294 (764)
PC Medicaid HH spend	$272 (281)	$244 (253)	$203 (220)	$198 (217)
Unemployment rate	5.08 (1.07)	4.98 (1.08)	5.04 (0.93)	4.97 (0.89)
% Population 65+	12.7 (1.9)	12.6 (1.1)	6.7 (1.0)	6.7 (1.0)
Tax on wage income				
Medicaid instrument				
N	36,451	3,511	23,492	17,722

Notes: Authors calculations from 1996 and 2001 SIPP panels. Standard errors in parentheses. All variables in dollars are real 2003 values. Samples of comparison workers in columns 3 and 4 are drawn from the first wave of each survey (March 1996 and January 2001).

Table 2: Detailed Summary of Wages

All Monthly Observations	
Mean	\$10.31
Standard deviation	\$6.08
Median	\$8.82
10^{th} Percentile	\$4.70
25^{th} Percentile	\$6.47
75^{th} Percentile	\$12.45
90^{th} Percentile	\$17.65
First Month of Spell	
Mean	\$9.49
Standard deviation	\$6.31
Median	\$7.96
10^{th} Percentile	\$3.65
25^{th} Percentile	\$5.62
75^{th} Percentile	\$11.38
90^{th} Percentile	\$17.16

Table 3: Detailed summary of employment spells (Length in months)

Non-censored (N=1,095)	
Mean	5.36
Standard Deviation	5.32
Median	4.0
10th Percentile	1.0
25th Percentile	2.0
75th Percentile	7.0
90th Percentile	12.0
Left-censored only (N=719)	
Mean	9.40
Standard Deviation	8.68
Median	6.0
10th Percentile	2.0
25th Percentile	4.0
75th Percentile	12.0
90th Percentile	23.0
Right-Censored Only (N=919)	
Mean	9.82
Standard Deviation	8.81
Median	7.0
10th Percentile	2.0
25th Percentile	3.0
75th Percentile	13.0
90th Percentile	23.0
Right and Left-Censored (N=890)	
Mean	15.76
Standard Deviation	14.0
Median	10.0
10th Percentile	3.0
25th Percentile	4.0
75th Percentile	27.0
90th Percentile	36.0

Table 4: Detailed Summary of Lifetime Months in Occupation

March 1996 (N = 532)	
Mean	112
Standard Deviation	98
Median	84
10th Percentile	12
25th Percentile	36
75th Percentile	168
90th Percentile	252
January 2001 (N= 669)	
Mean	126
Standard Deviation	111
Median	96
10th Percentile	12
25th Percentile	36
75th Percentile	180
90th Percentile	300

Table 5. Wage and Health Insurance Coverage Models

	(1) All wages	(2) 1st Month wages	(3) All ESI	(4) 1st month ESI
Age	0.0199	0.0218	0.0136	0.0141
	(3.31)**	(2.64)**	(1.95)*	(1.97)*
Age2	-0.0002	-0.0002	-0.0001	-0.0001
	(2.88)**	(2.39)*	(1.49)	(1.48)
Married	0.0467	0.0467	0.0454	0.0692
	(2.99)**	(2.11)*	(2.29)*	(3.42)**
Kids	-0.0146	-0.0398	-0.0864	-0.0926
	(0.78)	(1.55)	(3.39)**	(3.25)**
Hispanic	-0.0838	-0.0315	-0.0104	0.0276
	(3.49)**	(0.56)	(0.28)	(0.70)
Black	-0.0141	-0.0061	-0.0068	0.0129
	(0.58)	(0.20)	(0.22)	(0.51)
Other Non-White	-0.0171	-0.0014	0.0239	0.0165
	(0.42)	(0.03)	(0.52)	(0.35)
Metro area residence	0.0749	0.0342	0.0106	-0.0004
	(2.93)**	(1.29)	(0.45)	(0.02)
< HS diploma	-0.0862	-0.1033	-0.0956	-0.1456
	(4.33)**	(2.87)**	(2.73)**	(4.44)**
Some college	0.0409	-0.0156	0.0086	0.0004
	(1.69)*	(0.40)	(0.28)	(0.01)
Trade degree	0.1259	0.0522	0.0643	(0.01)
	(4.27)**	(1.07)	(1.97)*	(0.00)
Associate's degree	0.1231	0.1082	0.038	0.0534
	(2.69)**	(2.34)**	(1.17)	(1.28)
Bachelor's degree +	0.2962	0.2298	0.0659	0.0747
	(5.61)**	(4.26)**	(1.41)	(1.66)*
Union coverage	0.0491	0.0997	0.0151	-0.101
	(0.88)	(1.19)	(0.34)	(1.22)
Large firm	-0.0094	-0.0091	0.0579	0.0557
	(0.42)	(0.36)	(3.04)**	(3.37)**
Hospital	0.2359	0.1903	0.3165	0.2547
	(8.57)**	(4.75)**	(10.49)**	(11.07)**
Nursing home	0.0409	0.0383	0.1132	0.0581
	(1.47)	(1.19)	(4.64)**	(2.36)**
Community care	-0.0358	0.0294	0.0333	0.0074
	(0.88)	(0.49)	(1.00)	(0.26)
ln (% Manufact. employment)	0.0896	-0.4229	0.4391	0.2669
	(0.24)	(0.80)	(1.65)*	(0.79)
ln (% Retail employment)	-0.1943	0.1132	0.4663	0.0854
	(0.25)	(0.09)	(0.70)	(0.11)
ln (Minimum wage)	-0.2172	0.0454	-0.0234	0.1705
	(1.96)*	(0.30)	(0.29)	(1.33)
ln (PC Medicaid NH)	-0.0109	-0.0265	0.1415	0.0393
	(0.18)	(0.25)	(1.45)	(0.32)

ln (PC Medicaid HH)	0.0601	0.1127	-0.0534	0.0028
	(1.32)	(1.56)	(1.09)	(0.05)
ln Unemployment rate	-0.0085	0.006	0.002	-0.0176
	(0.47)	(0.20)	(0.11)	(0.74)
ln (% Population >65)	-0.961	-0.8758	-1.0094	-0.5205
	(1.75)*	(0.86)	(2.05)*	(0.65)
Wage pass-through	0.0699	0.0391	0.0105	0.004
	(3.06)*	(0.73)	(0.30)	(0.11)
Medicaid eligibility			-0.1464	-0.0435
			(1.49)	(0.42)
Income tax on wage			-0.0073	0.0367
			(0.28)	(0.75)

Notes: Marginal effects and absolute values of t-statistics are given. Sample size is 36,451 for Columns 1 and 3 and 3,505 for Columns 2 and 4. T-statistics are based upon robust Huber-White standard errors clustered by individual.

Table 7: Employment Duration Models

	(1) Non-LC	(2) LC	(3) Non-LC	(4) LC
ln (Wage)	0.78	0.76		
	(3.78)**	(5.08)**		
ESI	0.58	0.52		
	(7.86)**	(7.63)**		
ln (Starting wage)			0.90	0.85
			(2.46)**	(2.25)*
Starting ESI			0.62	0.70
			(6.50)**	(4.55)**
Age	0.99	0.99	0.99	0.99
	(3.20)**	(3.40)**	(3.02)**	(3.76)**
Child <5	1.26	1.05	1.26	1.09
	(2.03)*	(0.29)	(1.93)*	(0.55)
Children 6-10	1.16	1.03	1.14	1.03
	(1.11)	(0.17)	(0.91)	(0.20)
No children	1.12	0.93	1.11	0.93
	(1.61)	(0.73)	(1.44)	(0.81)
Married	0.91	1.07	0.91	1.06
	(1.35)	(0.80)	(1.33)	(0.71)
Hispanic	0.90	0.98	0.93	1.02
	(0.94)	(0.17)	(0.64)	(0.25)
Black	1.09	0.99	1.09	1.01
	(1.04)	(0.12)	(1.05)	(0.14)
Other non-white	1.13	0.88	1.09	0.86
	(0.60)	(0.63)	(0.39)	(0.72)
Metro area residence	1.14	1.29	1.12	1.25
	(1.43)	(2.18)*	(1.29)	(1.99)*
< HS diploma	1.18	1.01	1.20	1.05
	(1.33)	(0.10)	(1.45)	(0.48)
Some college	1.19	1.10	1.19	1.10
	(2.06)*	(0.85)	(2.06)*	(0.85)
Trade degree	1.39	1.53	1.36	1.46
	(3.34)**	(2.87)**	(3.31)**	(2.59)**
Associate's degree	0.92	1.14	0.88	1.11
	(0.73)	(0.59)	(0.99)	(0.49)
Bachelor's degree+	1.02	1.32	0.97	1.28
	(0.12)	(1.28)	(0.19)	(1.23)
Other income	1.01	1.01	1.01	1.01
	(1.59)	(1.40)	(1.38)	(1.01)
Union coverage	1.49	0.82	1.47	0.84
	(1.64)*	(0.56)	(1.52)	(0.50)
Large firm	0.89	0.87	0.88	0.86
	(1.69)*	(1.51)	(1.84)*	(1.61)
Hospital	0.89	0.91	0.84	0.81
	(1.12)	(0.86)	(1.62)	(2.04)*
Nursing home	1.05	1.10	1.03	1.07
	(0.51)	(1.01)	(0.33)	(0.69)
Community care	1.15	1.12	1.16	1.12
	(1.18)	(0.85)	(1.29)	(0.87)
ln (AFDC/TANF Ben.)	0.84	0.79	0.84	0.77
	(0.60)	(1.46)	(0.57)	(1.55)

ln (Minimum wage)	0.46	0.78	0.46	0.76
	(2.38)*	(0.57)	(2.50)*	(0.68)
ln (Unemployment rate)	0.63	1.29	0.64	1.33
	(1.20)	(0.41)	(1.23)	(0.48)
Secondary job	5.04	3.85	5.07	3.61
	(12.53)**	(11.95)**	(12.51)**	(11.27)**
Seam month	1.83	2.21	1.82	2.21
	(8.22)**	(9.85)**	(8.22)**	(10.08)**

Notes: Odds ratios and absolute z-values are given. Sample size is 36,451 for Columns 1 and 3 and 22,774 for Columns 2 and 4. T-statistics are based upon robust Huber-White standard errors clustered by state.

Appendix Table 1: State Long-Term Care Policy, 2002-2003

State	Wage pass through[1]	Minimum CNA training[2]	Medicaid nursing home $ per >65[3]	Medicaid home health $ per >65	Medicaid personal care[4]	State	Wage pass through	Minimum CNA training	Medicaid nursing home $ per >65	Medicaid home health $ per >65	Medicaid personal care
AL		75	$1020.59	$114.71							
AK		140	$1,409.88	$494.69	Yes	NE		76	$1,085.86	$64.63	Yes
AZ	Yes	120	$22.37	$47.53		NV		75	$497.19	$163.08	Yes
AR		75	$946.29	$111.02	Yes	NH		100	$750.76	$64.72	Yes
CA	Yes	150	$821.54	$698.57	Yes	NJ		90	$1930.10	$278.91	Yes
CO	Yes	75	$743.73	$246.40		NM		75	$635.39	$950.86	
CT		100	$2,432.55	$703.49		NY		100	$3,713.71	$1,208.95	Yes
DE		75	$1,393.10	$243.09		NC		75	$1,335.40	$309.76	Yes
DC		120	$3,459.15	$487.16	Yes	ND	Yes	75	$1,423.30	$21.40	
FL		120	$656.30	$118.13	Yes	OH		75	$2,251,01	$113.27	
GA		85	$1,012.84	$99.13		OK	Yes	75	$979.37	$91.00	Yes
HI		75	$543.71	$11.82		OR		150	$454.08	$59.52	Yes
ID		120	$1,060.55	$160.69		PA		75	$1,797.39	$85.26	
IL	Yes	120	$1,427.60	$37.03		RI	Yes	100	$1,375.84	$39.87	
IN		105	$1,425.71	$81.02		SC	Yes	80	$1,120.38	$25.41	
IA		75	$1,314.63	$113.21		SD		75	$752.76	$18.36	Yes
KS		90	$728.44	$96.00		TN		75	$1,350.78	$43.93	
KY		75	$1,087.27	$196.45		TX	Yes	75	$1,225.41	$424.35	Yes
LA	Yes	80	$1,949.62	$70.60		UT		80	$760.52	$73.13	Yes
ME	Yes	150	$1,381.63	$201.93	Yes	VT		75	$750.65	$212.69	
MD		100	$1,053.49	$332.01	Yes	VA	Yes	120	$958.39	$19.17	
MA	Yes	75	$1,678.67	$751.66	Yes	WA	Yes	85	$893.34	$281.88	Yes
MI	Yes	75	$1,032.41	$348.17	Yes	WV		120	$951.90	$108.17	Yes
MN	Yes	75	$1,457.22	$420.10	Yes	WI	Yes	75	$1,812.69	$401.10	Yes
MS		75	$1,708.76	$60.00		WY	Yes	75	$1,031.12	$66.52	
MO	Yes	175	$1,162.89	$280.56	Yes						

Sources: (1) Updated from Paraprofessional Healthcare Institute (2003) (2) Hernandez-Martina et al. (2006) and HHS (2002) (3) Centers for Medicare and Medicaid Services, National Health Expenditure Data (4) Kaiser Family Foundation (www.statehealthfacts.org)