Sex Differences in the Prevalence of Mobility Disability in Old Age: The Dynamics of Incidence, Recovery, and Mortality

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Objectives. This study examined sex differences in the prevalence of mobility disability in older adults according to the influences of three components of prevalence: disability incidence, recovery from disability, and mortality.

Methods. Participants in a population-based study of older adults from three communities in the United States (N = 10,263) were studied for up to 7 years. Life table methods were used to estimate the influence of each of the three components of disability prevalence in women and men. Sex differences in probabilities for transition states were measured by relative risks derived from a single model using a Markov chain approach.

Results. The proportion of disabled women increased from 22% of women aged 70 years to 81% of those aged 90 years. In men, comparable figures were 15% and 57%. Incidence had the greatest impact on the sex differences in disability prevalence until age 90 and older when recovery rates had a greater impact on differences in prevalence. Mortality differences in men and women had only a modest impact on sex differences in disability prevalence. These findings initially seemed to contradict striking sex differences observed in the relative risks for mortality in men compared with women. Subsequent graphical analyses showed that incidence rather than recovery or mortality largely accounted for sex differences in disability prevalence in old age.

Conclusion. Disability incidence, recovery from disability, and mortality dynamically influence the sex differences in the prevalence of mobility disability. However, incidence has the greatest impact overall on the higher prevalence of disability in women compared with men.

SEX differences in disability prevalence in older adults have been observed in numerous studies. The transitions that contribute to disability prevalence include disability incidence, recovery from disability, and mortality, each of which is influenced by aging, health status, and many other factors. Consequently, disability status is highly variable across individuals, and persons may move back and forth between states of nondisability and disability or from disability or nondisability to death. Although several studies have reported differences between men and women in each of these transitions, the magnitude of some of these differences varies between studies (Beckett et al., 1996; Pinsky, Leaverton, & Stokes, 1987; Strawbridge, Kaplan, Camacho, & Cohn, 1992; Verbrugge, 1976). Research has shown that women have a steeper rate of functional decline in old age compared with men across age and after controlling for age (Beckett, 1996; Crimmins, Saito, & Reynolds, 1997). Clearly, women have an advantage over men in terms of mortality, and men have an advantage over women in terms of duration of disability (Branch et al., 1991). However, it is unclear how the disability and mortality transitions interrelate to determine overall disability prevalence as people age.

Multistate life table methods have been used to account for the relative contribution of incident disability, recovery, and mortality to the overall prevalence of disability (Crimmins, Hayward, & Saito, 1994; Crimmins, Saito, & Reynolds, 1997). These approaches have contributed to an improved understanding of the relationship between the various transitions and have demonstrated how changing the rates of any one element can alter the overall prevalence (Crimmins, Hayward, & Saito, 1994; Crimmins et al., 1997). Nevertheless, explaining the gender gap in disability prevalence in terms of disability and mortality transitions remains a source of controversy. We know that women have lower mortality rates, which in turn contribute to more years of life at risk for disablement in the older ages. Related to their lower mortality, duration of disability is longer in women, particularly at the end of a long life (Branch et al., 1991; Crimmins, Land, Blazer, Fillenbaum, & Branch, 1993). Incidence of disability is generally higher in women, but adjustment for health and/or demographic factors in some studies diminished gender differences in disability incidence (Guralnik & Kaplan, 1989; Lawrence & Jette, 1996). An examination of the components of disability prevalence using multistate life table methods showed that mortality, perhaps more so than disability incidence, contributed to the longer disabled life expectancy in women compared with men (Crimmins, Hayward, & Saito, 1996; Hayward, Crimmins, & Saito, 1998). Others have suggested that attributing sex differences in disability prevalence to the later mortality in women does not account for differences in
recovery from disability as well as the steeper declines in functioning in women compared with men (Beckett et al., 1996). A demonstration of the dynamic interplay of the three factors would improve our understanding of how incidence, recovery, and mortality contribute to sex differentials in disability prevalence after age 65.

Another issue that kindles the debate about sex differences in disability prevalence is the assortment of definitions of disability, which in turn leads to a broad range of prevalence estimates from large population-based studies (Jette, 1994; Langlois et al., 1996). Often, disability is described as the inability to perform basic self-care activities, referred to as activities of daily living (ADL), without assistance. This is the most severe and least common form of disability, and, especially in the oldest old, ADL disability is most often the end result of a progressive disablement process (Ferrucci et al., 1996). A more prevalent form of disability is lower extremity mobility disability, often a precursor to further deterioration in functioning (Dunlop, Hughes, & Manheim, 1997; Lawrence & Jette, 1996). Because of the higher prevalence of mobility disability and the substantial differences in prevalence between women and men, mobility disability is an appropriate target for exploring sex differences in components of disability prevalence.

In this study, we sought to uncover whether any single disability or mortality transition, incidence, mortality, or recovery is key in determining sex differences in the prevalence of mobility disability according to age in older adults, or whether it is the dynamic interplay of the transitions and the accumulation of their effects over years of age that lead to the sex difference in disability prevalence. The Established Populations for the Epidemiologic Studies of the Elderly (EPESE), with its long study follow-up and annual assessments of disability status, is uniquely suited for a study of the dynamic influences on disability prevalence in older men and women.

**METHODS**

**Study Population**

The study population was older women and men from three communities of the EPESE: East Boston, Massachusetts; New Haven, Connecticut; and Washington and Iowa counties, Iowa. In East Boston and Iowa, the entire older adult populations living in the identified communities were invited to participate, and response rates were 85% and 80%, respectively. In New Haven, a stratified random sample was selected according to residence in private or public housing, and the oldest men were oversampled. The response rate in New Haven was 82% of eligible persons. Proxy informants were interviewed when participants were unable to answer the interview questions (8.8% of interviews in this analysis). In-home baseline interviews were conducted between 1981 and 1983, followed by annual interviews for 7 years in New Haven and Iowa, and 6 years in Boston. At the start of the study, there were 10,263 participants aged 65 to 95. Persons over age 95 (n = 31) were excluded from this study because of their small numbers. Mortality was determined by obituaries and proxy respondents and confirmed using death certificates. Details of the study design and methods were published previously (Cornoni-Huntley, Brock, Ostfeld, Taylor, & Wallace, 1986; Cornoni-Huntley et al., 1993).

**Measures**

Responses to two questions determined mobility status: “Are you able to walk up and down stairs to the second floor without help?” and “Are you able to walk half a mile without help? (about 8 blocks).” Persons who responded “no” to either of these questions were classified as having mobility disability. Those who responded that they were able to perform both activities without help were classified as nondisabled for the purposes of this study.

**Statistical Analysis**

**Data organization.—**To study disability transitions by year of age, each participant's history (which contained up to seven annual interviews) was split into pairs of consecutive interviews. Thus, a participant who completed all eight scheduled interviews had up to seven separate 1-year records, or consecutive interview pairs, given that there was complete disability and vital status information at each interview. If a person’s mobility status was known at the start and at the end of the time interval, or disability status was known at the start of the 1-year follow-up interval and she or he died before the next interview, the record was included in the study for that 1-year interval. Paired consecutive interview records containing missing functional status on persons not known to have died were omitted. There were 85 women (M age = 75.0) and 78 men (M age = 75.2) without complete information for at least two consecutive interviews. Among 10,100 participants, there were 52,286 observations with complete data for 1-year intervals during the study follow-up. More than half of participants had six or seven paired consecutive interviews and more than 75% of participants had at least four paired interviews.

**Disability transitions.—**In age- and sex-specific analyses, four measurements were used in describing the 1-year transitions in disability status: disability prevalence at the beginning of the year, disability incidence, recovery from disability, and mortality. Prevalence was measured as the proportion of persons of a specific age who were disabled at the start of the 1-year follow-up period. Incidence was measured as the proportion of those who were nondisabled at the start of the 1-year interval (age x) who reported being disabled at the follow-up interview (age x + n). Mortality during the 1-year interval was calculated separately among those who began the year as nondisabled or disabled in order to understand the separate influences of each rate on disability prevalence. In the disabled group, the mortality rate was described as the case fatality rate. The recovery rate was the proportion of disabled persons who reported being nondisabled at the end of the 1-year interval.

**Life table methods.—**Using the increment-decrement life table technique, we estimated the relative impact of each possible transition on the prevalence of mobility disability by age and sex. Beginning with 100,000 persons of each sex at age 65, we applied the disability prevalence at age 65 (from the EPESE) to initially distribute the population according to disabled and nondisabled states for each sex. The number of persons with each transition was then calculated sequentially by applying the transition probabilities (based on the proportions from the EPESE) for each consecutive year of age to the num-
numbers of persons remaining in the disabled and nondisabled groups from the end of the previous year. Using life table notation adapted from Crimmins et al. (1994), we calculated the proportion of persons with prevalent disability at a specific age $P_i(x + n)$ by the following:

$$P_i(x + n) = \frac{I_i(x) + d_{i0}(x,n) - d_{i1}(x,n) - d_{i2}(x,n)}{I_i(x) + I_2(x) - d_{i0}(x,n) - d_{i2}(x,n)},$$

where $x$ denotes age at the start of the 1-year interval and $x + n$ indicates the age at the end of the interval. Number of persons alive at a given age are represented by $I$. Number of persons who transitioned from one state to another are represented by $d$. The notation for state is as follows: $i = \text{disabled}$, $o = \text{nondisabled}$, $d = \text{death}$. For example, $d_{i0}(x,n)$ represents the number of persons who developed new disability during the 1-year interval, and $d_{i2}(x,n)$ is the number of persons who were not disabled at the start of the year and died during the 1-year interval. The numerator is the number of persons who are alive with disability at the end of the 1-year interval, and the denominator is all persons who are alive at the end of the interval. Through this method, we estimated the distribution of persons in each disability state at a given age on the basis of the transition probabilities and numbers of persons from the previous age.

The estimated prevalence of disability by age and sex served as a basis for evaluating the effects of each component of disability prevalence and thus a basis for understanding the impact of sex differences in incidence, mortality, and recovery on the overall prevalence. For example, the influence of recovery was examined by substituting the men’s recovery rates in the formula to calculate the women’s prevalence. This involved going back through the life table, using the women’s incidence rates and mortality rates and the men’s recovery rates, and recalculating the prevalence rates for each year of age. The true rates and the recalculated rates were plotted and examined graphically to detect visible differences between the women’s rates and the new scenario, of the women’s prevalence calculated with a substitution of the men’s recovery rates. Similarly, men’s incidence and mortality rates were also applied separately to the women’s data and again plotted and examined graphically.

Relative risks.—In addition to the descriptive approaches, a Markov chain model was used for making statistical inferences. The assumption in the Markov model is that each person’s likelihood for her or his given series of functional states over time could be expressed in terms of annual transition probabilities. This meant that the likelihood of any change in functional status during a single year interval was assumed to be independent of functional status in the previous years. Transition probabilities were modeled in a single model using logistic link functions for each of four dichotomous outcomes: death versus survival from the nondisabled state, death versus survival from the disabled state, and non-time-homogeneity for age groups. The Markov chain modeling was performed using a FORTRAN routine developed specifically for this purpose by Beckett and colleagues (Beckett, Brock, Scherr, & Mendes de Leon, 1993; Beckett et al., 1996; Mendes de Leon et al., 1997; Muenz & Rubinstein, 1985). The relative risks were derived according to sex and age group (5-year groupings). The method was modified to include time-dependent covariates for age groups.

For these analyses, each participant had a single record that included all annual interviews. All participants with disability and mortality information from at least two interviews were included in the analysis. Interviews with missing disability status flanked by interviews with nonmissing data were included in the likelihood by averaging the predicted probabilities of transitions for all possible paths between the first state and the next recorded state. This approach is based on the assumption that, conditional on the functional states immediately before and after the gap in the data, the data are missing at random. The justification for this assumption is discussed later. Of the 10,263 persons in the study, 57 people were excluded from this analysis because they had either no disability data or were missing data for all but one interview.

To determine whether time effects in the single year transition probabilities (non-time-homogeneity) were adequately modeled by age (e.g., ignoring the calendar effect of later versus earlier interviews in the study), separate models were fitted for each consecutive set of five interviews (interviews 1-5, 2-6, and 3-7). The results of this check revealed a reasonable amount of agreement between the age and sex effects to warrant the lack of calendar dependence within age groups.

The assumption of lack of dependence on functional states before the current state (Markovian dependence) was justified by the fact that the extra-Markovian residual correlation was extremely small. This correlation was estimated using Huber’s formula, treating all transitions within the individual as a cluster in the sampling. Within-person correlation was accounted for by using a robust variance estimate using the clusters, or treating each person’s data as a primary sampling unit. The assumption of the randomness of the missing data was also justified by the small size of the residual correlation. In addition, the distribution by age and sex of those who did not have disability information for at least two consecutive interviews (which was mentioned earlier in the Methods section) attests to the randomness of the missing information.

The stratified sampling by type of housing in the New Haven site was addressed in the robust variance estimate. This was done by scaling the original weights for the New Haven site to sum to the total number of New Haven participants and assigning a weight of 1 to each of the participants in the other two sites. Sex differences were measured in each component of prevalence according to mobility status at the start of each 1-year interval. Among the nondisabled, risk of incident disablement or death in one year was modeled for women compared with men within 5-year age groups. Similarly, in the disabled, likelihood of death or recovery was modeled in women compared with men. Adjustment variables in the models included age and EPESE site. Tests for linear interaction were conducted by adding an interaction term for sex and the continuous form of age to a model with main effects for age, sex, and EPESE site.

RESULTS

Prevalence and Transition Rates

Among the 10,100 EPESE participants, there were 32,642 single-year observations of women and 19,644 observations of men (Table 1). The proportions of persons disabled at the first interview in the 1-year observations increased steadily in both
sexes with age in all three EPESE sites. However, percentages of disabled women were consistently higher than percentages of disabled men within age groups by site and in the total cohort.

The age-specific prevalence of disability increased markedly from ages 65 to 95 in women and men, shown in Figure 1A. At age 70, the prevalence was 22% in women and 15% in men. By age 85, the sex differences were more pronounced, with a prevalence of 65% in women and 43% in men. The annual incidence of disability increased with age in both sexes (Figure 1B). At age 70, the incidence in women was 11% and in men was 7%. By age 85, the incidence in women was 33% and in men was 25%. Differences in the adults over age 85 were less clear because of the small numbers in the nondisabled cohorts of men and women.

Among persons with prevalent disability, sex differences and age trends were observed in both death rates (case fatality rate) and recovery rates. Men were more likely to die when disabled than were women, and the risk of death was higher with increasing age (Figure 1C). At age 70, the death rate for disabled men was 16% compared with 6% in disabled women and by age 85, 21% of disabled men and 12% of disabled women died within 1 year. Conversely, the annual rates of recovery from disability showed marked declines with advancing age (Figure 1D). At age 70, the recovery rate in women was 24% and in men was 29%. Among disabled women at age 85, only 10% recovered in the 1-year interval, and by age 90 the proportion was halved again. In disabled men aged 85, 13% recovered in 1 year and at age 90, 12% recovered, showing a low but stable rate of recovery in men aged 85 and older.

Comparisons of the likelihood of each outcome in women versus men across age groups revealed the relative impact of death, incident disability, and recovery on the probability of being disabled or not disabled at follow-up because each outcome directly influenced the others. Nondisabled women had approximately a 30 to 80% greater likelihood than men of becoming disabled during follow-up. However, in the oldest ages (90–95), the sex differences diminished (Table 2). Conversely, the risk of death in nondisabled women was approximately 2 to 3 times lower than that of nondisabled men across age groups, again with less of a sex difference in the oldest age group. There was no evidence of a linear interaction between age and sex for incident disability or death.

Among those who were disabled, the relative risk of death in women compared with men was generally similar to the relative risk of death in the nondisabled in most age groups, with women having up to a two-and-one-half-fold lower mortality risk (Table 2). Disabled women were less likely to recover compared with disabled men, and the differences were most extreme in the 90- to 95-year-old age group (test for Sex × Age interaction, \( p = .08 \)).

Sex Differences in Components of Prevalence

Graphical comparisons.—In order to understand the differences in disability prevalence based on age-specific rates of incidence, recovery and case fatality in women compared to men, we individually applied the men’s rates for each component of prevalence to the women’s data using life table methods (Figures 2A–2C). When we substituted men’s incidence rates for the women’s incidence rates in the formula for calculating women’s disability prevalence for each year of age, there was a
substantial shift to a lower prevalence rate until the oldest ages (90-95), when the sex differences widened (Figure 2A). Men’s recovery rates applied to the women’s data resulted in a gradual reduction in the sex differences with increasing age that became most pronounced after age 90 (Figure 2B), illustrating the modest Age X Sex interaction (shown in Table 2). Applying men’s case fatality rates to the women’s data had a moderate impact on prevalence estimates across ages until the oldest ages (90-95), when the effect diminished and the sex differences widened (Figure 2C).

In an additional analysis, overall death rates for nondisabled and disabled men were applied to the women’s data, and the results showed only a minimal change in the women’s disability prevalence (data not shown). Changes in death rates had a much greater impact on survival differences than on differences in disability prevalence. When the overall death rates for men in the nondisabled and the disabled groups were applied to the women’s data, only 16% of the cohort survived to 85 years of age compared with 44% using the women’s rates.

**Age-specific examples.**—The very low relative risks of women dying compared with men (see Table 2) do not at first seem compatible with the relatively greater impact that the sex differences in incidence have on disability compared with mortality (Figure 2). To illustrate the relative impact of the changes in transition rates across ages on the sex differences in prevalence, we calculated the distribution of single-year transitions in the entire cohort on the basis of a reference of 100 persons of each sex at three ages: 70, 80, and 90 years (Table 3). This simplified approach required substantial rounding but is used for descriptive purposes only. The more exact information is shown in the rates and in the rate ratios for each transition comparing women’s rates to men’s rates (using crude rates). At age 70, the greatest differences in rates between women and men, as shown by the rate ratios, were in the death rates among both the disabled and nondisabled groups. However, the numbers that died were small relative to the size of the cohort; thus, sex differences in death rates shown by the rate ratio had little impact on the numbers who were disabled or nondisabled at the end of the follow-up year.
Table 2. Relative Risks for Men Versus Women in Risk for Death, Disability, and Recovery From Disability by Age Group According to Disability Status

<table>
<thead>
<tr>
<th>Events</th>
<th>RR (95% CI)</th>
<th>RR (95% CI)</th>
<th>RR (95% CI)</th>
<th>RR (95% CI)</th>
<th>RR (95% CI)</th>
<th>RR (95% CI)</th>
<th>Interaction p value</th>
</tr>
</thead>
<tbody>
<tr>
<td>Nondisabled at age x</td>
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</tr>
<tr>
<td>Incident disability</td>
<td>1.46 (1.18—1.81)</td>
<td>1.41 (1.24—1.62)</td>
<td>1.36 (1.18—1.57)</td>
<td>1.56 (1.32—1.85)</td>
<td>1.80 (1.41—2.30)</td>
<td>1.04 (0.55—1.97)</td>
<td>0.82</td>
</tr>
<tr>
<td>Death nondisabled</td>
<td>0.37 (0.24—0.55)</td>
<td>0.41 (0.32—0.54)</td>
<td>0.32 (0.24—0.42)</td>
<td>0.35 (0.26—0.48)</td>
<td>0.43 (0.29—0.66)</td>
<td>0.73 (0.21—2.54)</td>
<td>0.32</td>
</tr>
<tr>
<td>Disabled at age x</td>
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<tr>
<td>Death disabled</td>
<td>0.80 (0.50—1.26)</td>
<td>0.43 (0.35—0.55)</td>
<td>0.44 (0.35—0.54)</td>
<td>0.43 (0.35—0.53)</td>
<td>0.54 (0.43—0.67)</td>
<td>0.45 (0.31—0.66)</td>
<td>0.78</td>
</tr>
<tr>
<td>Recovered</td>
<td>0.72 (0.53—0.98)</td>
<td>0.69 (0.56—0.85)</td>
<td>0.64 (0.51—0.78)</td>
<td>0.57 (0.45—0.73)</td>
<td>0.64 (0.47—0.88)</td>
<td>0.27 (0.13—0.58)</td>
<td>0.08</td>
</tr>
</tbody>
</table>

Note: The entire study sample aged 65 to 95 at baseline (N = 10,263), excluding 57 persons who did not have disability information from at least two interviews.

Relative differences between men and women's incidence rates were less profound, but incidence had a greater impact than mortality for two reasons. The incidence rates were somewhat higher than the death rates among the nondisabled and the pool of nondisabled persons was highest at the younger ages; thus the smaller differences in rates translated into larger num-

Figure 2. Mobility disability prevalence in men and women at the end of 1-year follow-up according to age. (A) Influence of men's incidence rates applied to women's data; (B) influence of men's recovery rates applied to women's data; and (C) influence of men's case fatality rates applied to women's data. Middle graph line shows disability prevalence in women based on substituting specified men's transition rates in the calculation of women's prevalence using life table methods. Data are for 10,100 EPESE participants from East Boston, Massachusetts; New Haven, Connecticut; and Iowa and Washington Counties, Iowa. 1982—1989.
bers of individuals. In other words, the rate ratio for deaths in the nondisabled adults showed a two-and-one-half-fold difference in rates between women and men, but in actuality this difference represented 1 more death in men compared with women based on a total of 100 persons in each group. However, a 50% higher incidence rate in women (rate ratio = 1.49) represented 2 more disabled women than disabled men entering the disability pool.

The reverse was true at age 90 when the differences between men's and women's recovery rates were highest (rate ratio = 0.46). At this age, there were many more people in the disabled pool; thus, even though recovery rates were low for both sexes, men had twice the recovery rate of women (11.6% vs. 5.3%, respectively). If one were to apply the men's recovery rates at age 90 to the women's data at the same age, the result would be 5 more women who recovered, for a total of 9 women, thus 5 more deaths among the disabled women. However, the denominator changed only in the latter example.

A more striking example was found in the application of the men's case fatality rate at age 80 (29.5%) to the women's data. This resulted in 24 deaths instead of 13 in the disabled women; however, the disability prevalence was reduced only to 81% (from 84%). Contrast this result with the previous example showing the impact of changes in recovery rates for the same age. Much smaller changes in the actual numbers (5 more women recovering from disability at age 90) resulted in a more notable reduction in prevalence (down to 78%). In summary, increases in the case fatality rates in women may result in the same or greater absolute change as increases in incidence rates, but changes in death rates among the disabled persons have a lesser impact on the overall prevalence than changes in incidence because the higher death rates reduce the total number of persons in both the numerator and the denominator of the prevalence calculation.

**DISCUSSION**

The profound differences in prevalence of disability in old age observed between men and women were due to the complex interplay of incidence, recovery, and death. In our study, disability incidence had the greatest influence on the overall differences in disability prevalence between men and women.

<table>
<thead>
<tr>
<th>Measure</th>
<th>Women N (Rate)</th>
<th>Men N (Rate)</th>
<th>Rate Ratio (W:M)</th>
<th>Women N (Rate)</th>
<th>Men N (Rate)</th>
<th>Rate Ratio (W:M)</th>
<th>Women N (Rate)</th>
<th>Men N (Rate)</th>
<th>Rate Ratio (W:M)</th>
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<tbody>
<tr>
<td>Among disabled</td>
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<tr>
<td>Recovery from disability</td>
<td>5 (23.5)</td>
<td>4 (29.2)</td>
<td>0.80</td>
<td>7 (16.1)</td>
<td>7 (19.8)</td>
<td>0.81</td>
<td>4 (5.3)</td>
<td>7 (11.6)</td>
<td>0.46</td>
</tr>
<tr>
<td>Died with disability</td>
<td>1 (5.8)</td>
<td>2 (16.1)</td>
<td>0.36</td>
<td>5 (10.6)</td>
<td>7 (21.8)</td>
<td>0.49</td>
<td>13 (16.2)</td>
<td>17 (29.5)</td>
<td>0.55</td>
</tr>
<tr>
<td>Remained disabled</td>
<td>16 (70.7)</td>
<td>9 (54.7)</td>
<td>1.29</td>
<td>32 (73.3)</td>
<td>20 (58.4)</td>
<td>1.26</td>
<td>64 (78.5)</td>
<td>33 (58.9)</td>
<td>1.33</td>
</tr>
<tr>
<td>Total disabled at age x</td>
<td>22</td>
<td>15</td>
<td></td>
<td>44</td>
<td>34</td>
<td></td>
<td>81</td>
<td>57</td>
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<tr>
<td>Among nondisabled</td>
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<tr>
<td>Incident disability</td>
<td>8 (10.6)</td>
<td>6 (7.1)</td>
<td>1.49</td>
<td>14 (24.2)</td>
<td>10 (15.5)</td>
<td>1.56</td>
<td>8 (41.6)</td>
<td>17 (38.9)</td>
<td>1.07</td>
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<tr>
<td>Died without disability</td>
<td>1 (0.9)</td>
<td>2 (2.5)</td>
<td>0.36</td>
<td>1 (1.9)</td>
<td>4 (6.6)</td>
<td>0.29</td>
<td>1 (5.2)</td>
<td>4 (8.3)</td>
<td>0.63</td>
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<tr>
<td>Remained nondisabled</td>
<td>69 (88.5)</td>
<td>77 (90.4)</td>
<td>0.98</td>
<td>41 (73.9)</td>
<td>52 (77.9)</td>
<td>0.95</td>
<td>10 (53.2)</td>
<td>22 (52.8)</td>
<td>1.01</td>
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<tr>
<td>Total nondisabled at age x</td>
<td>78</td>
<td>85</td>
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</table>

**Note:** Data are based on rates within EPESE cohorts (n = 10,100). Numbers were calculated on the basis of actual study rates (shown in parentheses). Differences between rates and numbers shown are due to rounding. Rates sum to 100% within categories of disabled and nondisabled persons.

*Ratio calculated as the women's rate divided by the men's rate.

*Age x indicates age at the start of the 1-year observation period.
abled had a lesser but steady impact on sex differences in disability prevalence until the oldest ages (90–95), and, beginning only after age 80, differences in recovery rates had an increasing impact on sex differences in disability prevalence. The net result was that women were more likely than men to become disabled and remain disabled, thus having longer duration of disability, particularly in the oldest ages.

The dramatic shift in the pools of nondisabled and disabled adults across ages contributed to the changes in the relative impact of incidence, recovery, and mortality on prevalence. From ages 65 to 95, the prevalence of disability changed from 16 to 90% in women and from 14 to 74% in men. Incidence had a strong impact on prevalence until very old age (age 90–95), when prevalence was high and few were nondisabled, thus incidence had a lessening absolute effect. Disabled men were consistently more likely to die than were disabled women. However, the effect of mortality on prevalence was less because it reduced the size of the entire cohort and not just the numbers of disabled persons. Recovery rates were higher in men than in women across ages, but the influence of recovery on the prevalence rates became most apparent in the later ages (≥90), when disability prevalence was highest, even though recovery rates declined markedly with age. Considering the combined effects of death and recovery on the disabled pool, another important sex difference was that men were less likely than women to remain disabled at the end of each follow-up year.

The gender differences in disability incidence shown graphically appear minor before age 80, such as a 3.5% difference in incidence rates between men and women at age 70 (7.1% vs 10.6%). In general, there was approximately a 5-year difference in incidence rates between men and women, such that women’s rates at any given age were about the same as rates in men who were 5 years older. Relatively small differences in incidence rates resulted in substantial differences in prevalence. The implications of the differences in incidence, in terms of relative risk, were that women had at least a 36% greater likelihood of developing disability between ages 65 and 89.

Longitudinal data from the Alameda County Study showed only slight sex differences in disability incidence and suggested that the sharp mortality differences between men and women, particularly in disabled persons, could explain the disability prevalence gap between men and women (Strawbridge et al., 1992). Analyses of the National Long-Term Care Survey (NLTCS) from 1982 and 1984 also supported the belief that sex differences in prevalence were related to differences in longevity rather than disability incidence (Manton, 1988). It is interesting that in the NLTCS, disability incidence was higher in men than in women at ages 65 to 84, yet greater mortality in men compared with women was observed in all age groups regardless of disability status. It is possible that differences in the study populations or in definitions of disability might account for the differences in these findings compared with our own. In our study, sex differences in mortality were greater than risk of incident disability in women versus men (shown in Table 2), thus we might also conclude that sex differences in mortality may have a greater impact than incidence on differences in overall prevalence. However, comparing the magnitude of the relative risks does not uncover the influence of the accelerating incidence rates with age that is shown using the life table approach. Use of the latter method and the substitution of men’s transition rates in calculating women’s disability prevalence revealed the greater importance of the sex differences in incidence. Application of similar methods to other data sets is needed to confirm that sex differences in disability incidence have a stronger influence than mortality on the differences in disability prevalence between men and women.

Previous studies have demonstrated that life table methods are uniquely suited to examinations of the relative impact of incidence, mortality, and recovery on prevalence. In studies of trends in disability, Crimmins and colleagues showed that disability incidence had a stronger effect on prevalence than either mortality or recovery (Crimmins et al., 1994, 1997). Using data from the Longitudinal Study of Aging (LSOA), Crimmins and colleagues showed that by separately applying the same percentage changes in mortality, incidence, and recovery to the calculation of disability prevalence, changes in incidence produced the greatest differences in the proportion of disabled persons in both sexes combined (Crimmins et al., 1997). Despite differences in the definitions of disablement, prevalence and mortality estimates in men and women in the LSOA were similar to the estimates in our own study, as were the relative risks for disablement and death in women compared with men (Crimmins et al., 1996, 1997). Although the purpose of our study and our use of life table methods differed from those of Crimmins and colleagues, each of the studies demonstrated the utility of life table methods in teasing out the relative impact of the components of prevalence.

When men’s mortality rates among the disabled group only were applied to the women’s data in our study, there was a modest shift in the curve of the women’s prevalence rates, reflecting the modest impact of case fatality on the gender differences in prevalence. It is interesting that when men’s death rates in both the disabled and nondisabled groups were applied to the women’s data, the sex differences were wider than when only case fatality rates were applied. In other words, applying men’s overall death rates for both the disabled and nondisabled groups to the women’s data had virtually no effect on narrowing the gender gap. It is possible that these dynamics are specific to mobility disability prevalence and may not operate the same way in prevalence of ADL disability. Crimmins et al. (1994) reported a modest effect on prevalence by changing the rates of overall mortality in her study of disability in ADL and in managing home activities such as shopping and cooking. Differences in definitions of disability (e.g., mobility vs ADL disability) may contribute to conflicting hypotheses on the relative contribution of mortality to disability prevalence. However, in both Crimmins’s work and our own, changes in incidence were shown to have the strongest effects on disability prevalence.

The findings about recovery from disability showed a substantial potential for reducing disability prevalence even in the oldest adults (age ≥ 85), when disability prevalence is very high. In the oldest ages, small changes in recovery rates could have a substantial impact on the proportion of people with disability. Our results were consistent with the physiologic reserve hypothesis (Buchner & Wagner, 1992) and suggest that men have greater reserves when faced with pathologic problems and are thus more likely to recover. Declines in muscle strength with aging pose a greater hazard for women, who have lower strength than men, in general. It has been shown that simple tasks such as stair climbing require the same level of strength in
both sexes, thus women have greater risk from mobility loss related to age-associated strength declines (Rantanen, Era, & Heikkinen, 1996). Greater attention to the maintenance of muscle strength in older adults has implications not only for disability incidence but also for recovery from disability.

The strengths of this study lie in the size of the study sample, drawn from diverse sites in the United States, the availability of annual interview data, and the long follow-up. Many previous studies of disability measurement and risk factors using EPESE data have attested to the validity and reliability of the disability information (Ferrucci, Guralnik, Pahor, Corti, & Havlik, 1997; Guralnik, LaCroix, et al., 1993; LaCroix, Guralnik, Berkman, Wallace, & Satterfield, 1993; Smith et al., 1990). The study populations were predominantly White, thus limiting the generalizability of the findings to similar populations.

There are potential biases inherent in self-reported disability information because it is possible that some of the gender differences may be related to reporting differences. Men may be less likely to report disability than women. However, mobility disability may be less prone to bias compared with other types of disability, such as those related to household activities or basic self-care activities (Johnson & Wolinsky, 1994). Recent comparisons between self-report disability and tests of physical performance, using data from the New Haven EPESE, found no sex differences in accuracy of self-reported mobility or ADL disability (Merrill, Seeman, Kasl, & Berkman, 1997).

The important message from this study is that incidence, recovery, and mortality have a dynamic influence on differences in mobility disability prevalence between older men and women, and the relative impact of each factor changes with age. However, the higher incidence of disability in women compared with men is the single most important factor leading to the sex differences in prevalence. The fact that women live longer than men has little bearing on the higher prevalence of mobility disability in women. Efforts to reduce the risk of developing disability or to postpone the development of disability are crucial to reducing disability prevalence in women and in men.

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