GENDER, BODY MASS, AND SOCIOECONOMIC STATUS: NEW EVIDENCE FROM THE PSID

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ABSTRACT

Previous research provides evidence of a negative effect of body mass on women's economic outcomes. We extend this research by using a much older sample of individuals from the Panel Study of Income Dynamics and by using a body mass measure that is lagged by 15 years instead of the traditional 7 years. One of the main contributions of this paper is a replication of previous research findings given our differing samples and measures. We compare OLS estimates with sibling fixed effects estimates and find that obesity is associated with an 18% reduction in women's wages, a 25% reduction in women's family income, and a 16% reduction in women's probability of marriage. These effects are robust – they persist much longer than previously understood and they persist across the life course, affecting older women as well as younger women.

From 1960 to 1999 the percentage of obese Americans nearly doubled so that today, more than one-third of all Americans are obese and more than two-thirds are overweight (Flegal, Carroll, Kuczmarski, & Johnson, 1998;

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Pastor, Makuc, Reuben, & Xia, 2002; Hedley et al., 2004). This is of concern because increased weight is associated with increased risk of morbidities and increased risk of mortality (Mokdad et al., 2001). It is also of concern because increased weight is associated with reduced economic well being, especially for women.

A number of recent studies provide evidence of a negative causal effect of body mass – a measure of weight for height – on U.S. individuals' economic and marital status outcomes (Averett & Korenman, 1996, 1999; Baum & Ford, 2004; Cawley, 2004; Cawley, Grabka, & Lillard, 2005). All but one of these studies (Cawley et al., 2005) use data from the National Longitudinal Survey of Youth (NLSY) where respondents are at relatively early stages in their lives. Two other studies, published in the current issue, provide evidence of a negative causal effect of body mass on older individuals' economic outcomes, but these studies focus on the European context (Garcia & Quintana-Domeque, current issue; Lundborg, Bolin, Hojgard, & Lindgren, current issue). In contrast, we build on this research but analyze data from the Panel Study of Income Dynamics (PSID) where U.S. respondents are, on average, 20 years older than NLSY respondents. In addition, we employ a body mass measure that is lagged by 15 years instead of the traditional 7 years. Despite these two substantial changes in our methods relative to previous studies, we report similar results: increased body mass is associated with a reduction in women's – particularly White women's – hourly wages, family income, and probability of marriage. Our findings attest to the robustness of the body mass penalty for women – it persists much longer than previously thought, and it persists across the life course, affecting younger women as well as older women.

1. PREVIOUS RESEARCH ON BODY MASS AND SOCIOECONOMIC STATUS

Most studies have found that body mass is negatively correlated with women's socioeconomic status and is insignificantly, or minimally, correlated with men's socioeconomic status. Register and Williams (1990) used data from the NLSY to compare wage rates and obesity among individuals who were between the ages of 18 and 25 in 1982. Even after correcting for a number of measured factors that affect wage rates – such as union status, race, work experience, and industrial category – the authors found that obese women earn 12% less than their non-obese counterparts. They found

that obese men, however, earn only 5% less than their non-obese counterparts. Pagan and Davila (1997) also found that women pay an earnings penalty for obesity and that men do not, primarily because men are able to sort themselves into occupations which may offset potential obesity penalties.

Cross-sectional associations, however, may be biased. As Cawley (2004) describes in detail, the negative association between women's economic status and their body mass could reflect one of the three processes: (1) diminished socioeconomic status could cause an increase to body mass through mechanisms such as reduced mental health or reduced quality food options; (2) increased body mass could directly reduce socioeconomic status through mechanisms such as employer discrimination, reduced health, or reduced productivity; or (3) cross-sectional associations could be caused by unobservable characteristics that are correlated with both socioeconomic status and body mass. To disentangle these three processes, researchers have employed a number of techniques which include the use of multiple waves of data with lagged body mass measures; sibling or individual fixed effects models; or most recently, instrumental variable (IV) models.

Gortmaker, Must, Perrin, Sobol, and Dietz (1993) use panel data from the NLSY to examine the correlation between obesity and several socioeconomic outcomes including education level, household income, marriage and poverty probabilities, the probability of graduating from college, and self-esteem. With the exception of self-esteem, the authors find that women who were obese between the ages of 16 and 24 suffer from worse outcomes 7 years later, at ages of 23–31. When a number of background measures are held constant, differences in marital status, income, poverty rates, and years of schooling remain significant for women. The only significant negative effect of obesity for men is a reduction in their likelihood of marriage.

Other recent studies have paired a lagged measure of body mass with sibling or individual fixed effects to control for unobserved family- or i ndividual-level heterogeneity. Behrman and Rosenzweig (2001) use data from 402 monozygotic twin pairs from Minnesota and find that negative cross-sectional associations between body mass and wages are reflective of unmeasured characteristics. Controlling for these "earnings endowments" with twins reduces the association to zero. However, twin studies are of questionable generalizability to the U.S. population on the whole (see, for example, Bound & Solon, 1999).

Averett and Korenman (1996) use the NLSY and employ sibling fixed effects models as a control on unmeasured family background characteristics. The authors examine wage rate differentials as a more direct test of

discrimination, and they break down income differentials into a component related to the job market – wages – and into a component related to the marriage market – the likelihood of being married and spouse's earnings. As with other studies, the authors report negative correlations for women, but mild or insignificant correlations for men. Further, these negative correlations are most pronounced for White women, and are primarily concentrated in the marriage market. Differences in marriage probabilities between overweight and healthy weight women account for 50–95% of their lower socioeconomic status. However, sister fixed effects wipe out the correlations in obesity and marriage and economic outcomes for all models except the one which predicts total family income without controls for marital status, children, and age of the youngest child.

In a subsequent analysis, Averett and Korenman (1999) use the NLSY to examine mechanisms that might account for Black–White differences in the effect of obesity on women's economic statuses. They evaluate economic status in 1990 when individuals are between the ages of 25 and 33, and they evaluate body mass in 1982 when individuals are between the ages of 17 and 24. The authors find that self-esteem does not explain racial differences in the obesity penalty. The negative effects of obesity for White women work largely through the marriage market as obesity significantly reduces White women's likelihood of marriage and significantly reduces their spouse's earnings.

Cawley (2004) also uses a relatively young sample from the NLSY and employs an individual fixed effects model as well as an IV model. Individual fixed effects models, which control for all time-invariant unobservable variables specific to individuals, reveal no significant negative effect of body mass or body weight on Black women's, Hispanic women's, or Hispanic men's hourly wages. These fixed effects models, however, reveal a significant negative effect of body mass on White women's hourly wages. Using the instrument of sibling's body mass as a putatively exogenous source of variation in an individual's body mass, Cawley reports an estimate of body mass on wages that is also only significant for White women. Cawley et al. (2005) extend this research with a comparison of the obesity penalty between German and U.S. individuals. The authors use the PSID and the sibling body mass instrument to model the relationship between body mass and earnings in the U.S., and they find that in both the IV and the ordinary least squares (OLS) models, body mass causes a significant reduction to U.S. women's earnings.

However, it is possible that the instrument of sibling's body mass fails to meet the exclusion assumption necessary for producing unbiased estimates,

as it may be directly correlated with an individual's socioeconomic status. Having an overweight sibling may lead you down certain economic and employment paths. Further, the IV approach does not deal with the problem of unobserved heterogeneity, since the same lurking variable that causes the sibling's body mass to be associated with the respondent's in the first stage regression may be associated with the error term in the second stage. But the fact that Cawley's (2004) and Cawley et al.'s (2005) results are consistent across OLS and IV models (and with corrections for measurement error) makes these studies the most definitive statements on the casual relationship between body mass and economic status to date.

Lundborg et al. (current issue) also use an IV approach in their study of the effect of obesity on European men's and women's labor supply and wage rates. Their European results echo what others have found for the U.S. population – that obesity has no effect on men's hourly wages but serves to reduce women's hourly wages by about 10%. Upon the inclusion of health status variables in their models, women's obesity wage penalty is reduced by only 1 percentage point, which suggests that the obesity penalty may derive from employer-related discrimination and not from reduced productivity due to diminished health status – to the extent that health status is measured without error. In their wage models, the authors induce exogenous variation in obesity status through the instrument of whether a female respondent has only sisters and no brothers. Lundborg et al. note that this instrument may not satisfy the exclusion assumption as it may be correlated with the error in predicting women's wages, and with this strategy they are unable to reject the hypothesis of obesity as exogenous to women's wages. Garcia and Quintana-Domeque (current issue) also provide evidence of a negative correlation between obesity and European women's wage rates. Though the authors provide a descriptive account only, their estimates of a 7–10% obesity wage penalty, depending upon which European country is under study, are very similar to Lundborg et al.'s estimates.

2. THE CURRENT STUDY

While previous studies have used various methods to reduce bias and to approach causal estimates of the effects of U.S. individuals' body mass on their socioeconomic statuses, all but one of these studies (namely Cawley et al., 2005) are limited in their reliance on samples from the NLSY where individuals are at early stages in their lives. In contrast, we analyze a different survey, the PSID. The PSID obtained respondents' height and weight in

1986 and during the latest two survey waves available, 1999 and 2001. We compare siblings from these three survey years who also have valid data available on their education level, wages and earnings (if employed), marital statuses, and total family income. These respondents are, on average, much older than respondents from the NLSY samples previously used, and they allow for a comparison between younger and older individuals, or those at different stages in the life course.

Our minimum age for inclusion into the sample is 25 in 1986, the year in which we first measure body mass. In 1986 women and men were, on average, 49 and 46 years old, respectively. In that same year, NLSY respondents were between the ages of 21 and 29. We measure economic and marital status outcomes 13–15 years later when respondents are, on average, around 60 years old (see Table 1). Unlike data from the NLSY, our data on marital status and economic outcomes *follow* the completion of formal education, and the peak years of earnings and childbearing. We examine body mass within OLS and sibling fixed effects models for marital status outcomes and for three economic outcomes – hourly wages, labor market earnings, and total family income.

3. DATA AND METHODS

The PSID began in 1968 with a nationally representative sample of 5,000 American families and has followed them each year since. It is a complicated study design and cannot be done justice in the space allowed here. For a fuller description, see Hill (1992) or Duncan and Hill (1989). By virtue of its complex design, this study has information on the socioeconomic histories of families, as well as on the outcomes of multiple children from the same families who were in the original sample, moved into it, or were born to sample members. We select adult respondents aged 25 and older who were head or wife of their household in any (or all) years for 1986, 1999, and 2001. Further, these individuals had to have a valid person number for their mother; that is, their mother had to have been in the sample at some time. They were then linked to their siblings through this maternal connection. A trivial number (less than 1%) of respondents had a father in the sample but not a mother. The majority had both parents. But since many more of the fathers were missing, we decided to identify siblings based on their mother's identification.

We describe our body mass measures and the variables that we use to capture socioeconomic status below. For sibling fixed effects models, the

Table 1. Sample Means by Gender and Body Mass (Standard Errors).

	Women				Men			
	Healthy Weight	Underweight	Overweight	Obese	Healthy Weight	Underweight	Overweight	Obese
Outcome Variables								
Log wages	2.58 (0.03)	2.47 (0.10)	2.43 (0.07)	2.34 (0.07)	3.04 (0.04)	3.37 (0.24)	3.00 (0.04)	2.77 (0.09)
Log earnings	9.82 (0.05)	9.52 (0.19)	9.68 (0.11)	9.55 (0.15)	10.58 (0.05)	11.19 (0.28)	10.58 (0.06)	10.19 (0.15)
Log family income	10.78 (0.04)	10.86 (0.13)	10.25 (0.09)	10.22 (0.12)	11.04 (0.04)	10.90 (0.13)	11.01 (0.05)	10.65 (0.12)
Proportion married	0.67 (0.02)	0.71 (0.06)	0.50 (0.03)	0.46 (0.04)	0.78 (0.02)	1.00 (0.00)	0.80 (0.02)	0.75 (0.04)
Proportion divorced	0.29 (0.02)	0.27 (0.06)	0.46 (0.03)	0.48 (0.04)	0.17 (0.02)	0.00 (0.00)	0.18 (0.02)	0.23 (0.04)
1999 Control Variables								
Labor market experience	2.91 (0.06)	2.83 (0.18)	2.71 (0.11)	2.77 (0.16)	4.98 (0.07)	4.91 (0.85)	5.42 (0.08)	5.20 (0.15)
Age	57.76 (0.48)	50.21 (1.64)	63.54 (1.00)	59.60 (1.06)	55.68 (0.61)	57.60 (7.91)	58.01 (0.59)	57.40 (1.05)
Children under the age of 1	0.01 (0.00)	0.04 (0.02)	0.01 (0.01)	0.01 (0.01)	0.02 (0.01)	0.20 (0.20)	0.01 0.00	0.01 (0.01)
Children under the age of 18	0.24 (0.01)	0.42 (0.07)	0.19 (0.03)	0.14 (0.03)	0.31 (0.02)	0.02 (0.02)	0.01 (0.00)	0.25 (0.04)
Educational attainment	13.22 (0.08)	13.35 (0.32)	11.95 (0.18)	11.81 (0.22)	13.63 (0.11)	14.20 (1.59)	13.30 (0.13)	12.32 (0.27)
Black	0.05 (0.01)	0.06 (0.03)	0.11 (0.02)	0.18 (0.03)	0.04 (0.01)	0.11 (0.11)	0.06 (0.01)	0.09 (0.02)
Number of observations	1,143	96	343	206	742	10	617	183

unit of analysis is the sibling and we have averaged economic outcomes from 1999 and 2001. Mean values – which generally conform to national averages – are presented in Table 1.

3.1. Economic Outcomes

The three economic outcome variables that we use in this study are hourly wage rates, labor market earnings, and total family income. All three variables are logged to the base e to correct for skewness.

3.2. Marital Status Outcomes

We examine the effects of body mass on the likelihood of being married or being divorced, separated, or widowed in 2001. In this year, about 62% of women are married and about 78% of men are married. On average, healthy weight women are more likely to be married than overweight or obese women, but healthy weight men are no more likely to be married than overweight or obese men. The same holds true for divorce. About 29% of healthy weight women are divorced in 2001, whereas about 46% of overweight women are divorced and 48% of obese women are divorced. The proportion of divorced men does not vary by body mass – about 17% of healthy men are divorced in 2001, whereas about 18 and 23% of overweight and obese men are divorced in 2001.

3.3. Body Mass

Body mass index (BMI) is a ratio of weight (in kilograms) to height (in meters squared). The PSID collects weight (in pounds) and height (in inches). We converted these measures but excluded all extreme values (weight greater than 400 lb or less than 70 lb and height greater than 84 in. or less than 45 in.). In most models, we compare obese, overweight, and underweight individuals to healthy weight individuals. To determine these categorical statuses, we rely on the clinical classifications where an individual is obese if her BMI is at or above 30, overweight if her BMI is at or above 25 but below 30, underweight if her BMI is below 18.5, and at a healthy weight if her BMI is at or above 18.5 but below 25. Though our primary 1986 predictor variables are obese, overweight, or underweight

compared with healthy weight, we use a continuous measure of BMI logged to the base e in some models.[1] In 1986, women had an average body mass of 24.20 and men had an average body mass of 25.82.

3.4. Control Variables

All models control for respondent's age (models are robust to the inclusion of a quadratic age term as well). Models also control for educational attainment. This is measured as the total number of years of formal schooling completed – a continuous variable from 1 to 17, with the topcode representing any graduate work, regardless of whether a degree was received. (The PSID does not, unfortunately, distinguish between various levels of graduate schooling.) Models also control for labor market experience, which is measured using PSID data on weekly work hours for all years from 1968 to 2001. Additionally, models control for respondent's parenthood status with three dummy variables for no children living in the household, children under age 1 living in the household, and children between the ages of 2 and 18 living in the household. Coefficients remain relatively robust to these control variables and robust to the inclusion of either lagged 1986 control variables or non-lagged 1999–2001 contemporaneous variables.

3.5. Statistical Approach

To estimate the effect of body mass on socioeconomic status, we would ideally like to control for all factors that affect body mass. However, there are likely a host of genetic and environmental factors that are unmeasured and unobserved in the PSID (and in any survey sample). To the extent that these unobserved factors are associated with body mass and with socioeconomic and marital statuses, OLS estimates would produce inconsistent results. For example, if we take the case of parent's body mass – an unobserved factor – that is positively associated with an individual's body mass and negatively associated with an individual's socioeconomic status, then OLS would produce upwardly biased estimates of a body mass effect. It is unclear if unobserved variables serve only to produce upwardly biased body mass estimates or if there are unobservable variables which may also lead to downwardly biased estimates. For example, certain parents may foster a sedative learning environment for their children and this unobserved factor would be positively associated with an individual's body mass and

also positively associated with their socioeconomic status. This would lead to a downwardly biased OLS estimate.

As discussed above, researchers have used sibling comparisons to deal with the issue of unobservable variable bias, as siblings are more alike on genetic and environmental factors than are a random sample of individuals. Differencing estimates across siblings effectively reduces many of the unobserved, confounding environmental and genetic factors. The sibling fixed effects model is specified in Eq. (1):

$$Y_{ij} = \beta \mathbf{X}_{ij} + \alpha_i + \mu_{ij} \tag{1}$$

where Y_{ij} is our outcome variable of interest for sibling j in family i (namely socioeconomic and marital status outcomes), $\mathbf{X}$ is a vector of body mass and explanatory variables, and our error term is broken into two components: α_i, the family fixed effect, and μ_{ij}, the error specific to each sibling j in family i. When we difference across siblings in each family, the unobserved family fixed effect is eliminated. This method does not, however, eliminate factors that are specific to each sibling j in family i. To the extent that these non-sibling-constant errors are correlated with the explanatory variables, the estimates from sibling fixed effects models may still be biased.

Sibling fixed effects models capitalize only on the variation in explanatory and outcome variables that occurs *within* families and *between* siblings. As such, they return strict and inefficient estimators, and variation between families is lost. Therefore, we compare sibling fixed effects models with more efficient OLS models which analyze variation between families. For all models, we report results from a Hausman (1978) test of the hypothesis of no systematic differences between OLS and fixed effects estimates. In most of our models, we find few systematic differences between OLS and fixed effects coefficients.

Ideally, we would like to estimate the causal effect of body mass on socioeconomic and marital status outcomes without bias due to unobserved family heterogeneity (as discussed above) *and* without bias due to a specific form of endogeneity – that of reverse causality. Sibling fixed effects models alone do not effectively eliminate bias due to reverse causality. In fact, it is likely that reverse causality is especially problematic in the current study as socioeconomic and marital status outcomes could be significant causes of body mass. We follow the lead of previous research that uses a lagged body mass variable to deal with bias due to reverse causality. Instead of using a lag of 7 years – as do Averett and Korenman (1996), Cawley (2004), and Gortmaker et al. (1993) – we use a lag of 13–15 years. We examine the effect of 1986 body mass on averaged 1999–2001 socioeconomic outcomes and on

2001 marital status outcomes. Though our use of a 15 year lag is largely by default – the PSID only includes weight and height variables for the years 1986, 1999, and 2001 – it nevertheless provides a substantively important test of the persistence of the obesity penalty across a much longer time period than has been previously examined.

4. FINDINGS

We begin with an initial analysis of the extent to which socioeconomic status, height, weight, and body mass cluster within families. To estimate sibling resemblances, we use a variance decomposition method that follows the strategy for income used by Mazumder and Levine (2003) and Solon, Corcoran, Gordon, and Laren (1991). See Conley and Glauber (2005) for a thorough discussion of this variance decomposition method for an unbalanced survey design.

Sibling correlations for height, weight, and body mass are calculated for same-sex siblings and are presented in Table 2. We report a 0.359 correlation in sisters' height, a 0.292 correlation in sisters' weight, and a 0.332 correlation in body mass. For brothers, we find a 0.529 correlation in height, a 0.371 correlation in weight, and a 0.318 correlation in body mass.

These figures imply that there is a significant amount of sibling variation in body mass – even more than the amount of intra-family variation in most

Table 2. PSID Sibling Correlations in Socioeconomic Status Using 1983–2001 Waves and Sibling Correlations in Height, Weight, and Body Mass for Sisters and Brothers Using 1986, 1999, and 2001 Waves (Fisher's z transformation, Standard Errors of z, Number of Person-Years, Number of Individuals, and Number of Families Below).

Ln Earnings	Ln Income	Height: Sisters Only	Height: Brothers Only	Weight: Sisters Only	Weight: Brothers Only	BMI: Sisters Only	BMI: Brothers Only
0.376**	0.458**	0.359**	0.529**	0.292**	0.371**	0.332**	0.318**
0.395	0.495	0.376	0.589	0.301	0.389	0.345	0.330
0.035	0.035	0.023	0.026	0.023	0.026	0.023	0.026
20,792	18,144	5,042	3,926	4,949	4,385	4,873	3,916
1,876	1,871	2,188	1,849	2,171	1,904	2,165	1,849
801	806	1,847	1,505	1,833	1,536	1,830	1,505

** $p < 0.01$.

socioeconomic measures. The degree of variation on these indicators within families is relevant to the current study because our sibling fixed effects models capitalize only on variation that occurs within families and between siblings.

5. THE BODY MASS SOCIOECONOMIC STATUS PENALTIES BY RACE AND GENDER

Table 3 presents the effect of obesity, overweight status, and underweight status in 1999–2001 on average log hourly wages, log labor market earnings, and log family income. We report both OLS and sibling fixed effects models and include in all of our models educational attainment, labor market experience, age of youngest child, and age. We include race only in the OLS models, as race does not vary between siblings. The bottom two rows in Table 3 present the chi-square test statistic of the Hausman test and the probability of obtaining this test statistic. Where the Hausman test leads us to reject the null hypothesis of no systematic differences between OLS and sibling fixed effects models, we report fixed effects estimates. In most models, there are few systematic differences between OLS and fixed effects estimates.

In line with previous research, we find that obesity is associated with a 17.51% reduction in women's wages and a 25.06% reduction in women's family income. Overweight status, as compared with healthy weight status, causes a 20.87% reduction in women's family income. The negative association between women's body mass and labor market earnings may be due to unobserved heterogeneity as sibling fixed effects models do not result in significant point estimates. Table 3 also reveals that for men, being obese or overweight does not come with any economic penalties.

Table 4 reports estimates from models which include a continuous measure of body mass and an interaction effect between race and body mass. The continuous measure of body mass is necessary for our sibling fixed effects models which rely on inter-sibling differences. Our results from Table 4 parallel results presented in Table 3 – body mass negatively affects women's wages, earnings, and family income, although the standard errors are higher for the wage and earnings models which result in statistically insignificant estimates. These findings also indicate that body mass penalties are amassed by White women and not by African American women. Black–White differences are most pronounced for family income. For women's family income, the OLS estimates do not appear to be biased by unobserved heterogeneity. The coefficients for wages and earnings work in similar

Table 3. OLS and Sibling Fixed Effects Models Predicting 1999–2001 Log Wages, Log Earnings, and Log Family Income from 1986 Body Mass (Standard Errors).

	Women						Men					
	Log Wages		Log Earnings		Log Family Income		Log Wages		Log Earnings		Log Family Income	
	OLS	Sib FE	OLS	Sib FE	OLS	Sib FE	OLS	Sib FE	OLS	Sib FE	OLS	Sib FE
Obese	−17.51**	−27.30	−25.16*	−40.46	−25.06*	−4.29	−13.42	−10.42	−24.41	−19.75	−18.82	−43.68
	(0.07)	(0.24)	(0.14)	(0.37)	(0.13)	(0.26)	(0.09)	(0.20)	(0.14)	(0.20)	(0.13)	(0.41)
Overweight	2.79	23.84	12.07	72.21*	−20.87*	−17.14	−2.10	15.35	0.76	22.62	3.89	6.56
	(0.07)	(0.18)	(0.11)	(0.27)	(0.09)	(0.19)	(0.05)	(0.12)	(0.07)	(0.11)	(0.06)	(0.23)
Underweight	0.00	0.00	−0.40	−4.69	−0.84	−3.32	−0.84	−5.62	−1.12	−0.98	−1.27	−15.97
	(0.09)	(0.24)	(0.17)	(0.37)	(0.14)	(0.30)	(0.21)	(0.83)	(0.25)	(0.82)	(0.25)	(1.81)
Number of individuals	623	623	630	630	1,052	1,052	712	712	719	719	1,023	1,023
Number of siblings		553		560		970		592		600		884
Chi-square test of difference in OLS and FE is not systematic		15.39		41.10		11.20		7.20		14.13		10.10
Probability > chi-square		0.052		0.000		0.191		0.515		0.078		0.258

Notes: All models include control variables for 1999–2001 educational attainment, labor market experience, age of youngest child, and age. Race is included in OLS models but not in sibling fixed effects models because it does not vary between siblings. All models report the percent change in wages, earnings, or family income due to obesity, overweight, or underweight as compared with recommended weight. To obtain this percent change, we used the following transformation: 100*(exp(B−S.E.*S.E./2)−1). Robust standard errors are reported for OLS models.

*$p<0.05$.

**$p<0.01$.

Table 4. OLS and Sibling Fixed Effects Models Predicting 1999–2001 Log Wages, Log Earnings, and Log Family Income from a Continuous Measure of 1986 Log Body Mass, Race, and Interactions (Standard Errors).

	Women						Men					
	Log Wages		Log Earnings		Log Family Income		Log Wages		Log Earnings		Log Family Income	
	OLS	Sib FE	OLS	Sib FE	OLS	Sib FE	OLS	Sib FE	OLS	Sib FE	OLS	Sib FE
Log body mass	−0.27**	−0.56	−0.35	−0.17	−0.71***	−0.44	−0.17	0.52	−0.43	−0.12	−0.18	−0.16
	(0.12)	(0.43)	(0.22)	(0.68)	(0.22)	(0.45)	(0.24)	(0.40)	(0.28)	(0.39)	(0.26)	(0.80)
Black	−0.10	0.00	−0.16	0.00	−4.67**	0.00	3.70*	0.00	0.37	0.00	0.62	0.00
	(1.33)	(0.00)	(1.88)	(0.00)	(2.08)	(0.00)	(2.22)	(0.00)	(2.86)	(0.00)	(1.95)	(0.00)
Black × log body mass	−0.00	0.75	0.04	1.43	1.36**	0.53	−1.24*	−4.09	−0.16	−0.92	−0.21	−4.56
	(0.41)	(0.95)	(0.59)	(1.49)	(0.63)	(1.06)	(0.68)	(4.31)	(0.88)	(4.29)	(0.59)	(9.57)
Number of individuals	624	624	631	631	1,053	1,053	713	713	720	720	1,024	1,024
Number of siblings		554		561		971		592		600		884
Chi-square test of difference in OLS and FE is not systematic		15.68		31.96		10.15		7.31		13.92		10.07
Probability > chi-square		0.028		0.000		0.184		0.605		0.125		0.345

Notes: All models include control variables for 1999–2001 educational attainment, labor market experience, age of youngest child, and age. Robust standard errors are reported for OLS models.

*$p<0.10$.

**$p<0.05$.

***$p<0.01$.

directions, although their standard errors are higher which result in insignificant estimates.

We are not the first to report race differences in the effects of body mass on economic outcomes. Averett and Korenman (1996) also find that the negative effects of obesity are more pronounced for White women than those for African American women. In fact, they find that overweight Black women earn about 8% more per hour than obese White women. We report similar results although we find only significant positive effects of body mass on African American women's family income. For a variety of social, political, and historical reasons, there seem to be fewer social and economic penalties for overweight Black women. However, we do not take this to mean that Black women's life outcomes are somehow less dependent on health or physical attractiveness. Instead, a comparably important measure of beauty for Black women is that of their skin color, as lighter skinned African Americans are still more likely to be considered physically attractive than are their darker peers and enjoy higher SES (Hughes & Hertel, 1990).

For men, we find that body mass has no deleterious economic effects, but with one exception. African American men pay a wage penalty for their body mass. The same does not hold for White men.

Findings reported in Table 5 indicate that body mass reduces women's likelihood of marriage and increases women's likelihood of divorce, separation, or widowhood. At the sample means of explanatory variables, obesity leads to a 0.16 reduction in women's probability of marriage, and overweight status leads to a 0.12 reduction in women's probability of marriage. For the probability of marriage, body mass OLS coefficients are not biased due to sibling-constant unobservable variables. For divorce, however, unobserved sibling heterogeneity produces upwardly biased OLS coefficients. The OLS model indicates that on average, obesity leads to a 0.13 increase in the probability of divorce and overweight status leads to a 0.11 increase in the probability of divorce, but estimates from sibling fixed effects models are not significantly different from zero.

In summary, findings that we have presented thus far replicate results from previous studies. Women, White women in particular, pay economic and marital status penalties for being obese and overweight, and men do not. Given that our samples and measures differ substantially from previous studies, our findings attest to the robustness of the body mass penalty for women as it persists across the life course. We next turn to an analysis of variation in the body mass penalty between older and younger women. Here we capitalize on the older age structure of the PSID as compared with the NLSY and present new research findings.

Table 5. Marginal Effects of Women's 1986 Body Mass on Women's 2001 Marital Statuses from Logit and Logit Sibling Fixed Effects Models (Standard Errors).

	Married		Divorced	
	Logit	Sib FE	Logit	Sib FE
Obese	−0.16**	−0.09	0.13*	0.07
	(0.05)	(0.14)	(0.05)	(0.16)
Overweight	−0.12**	−0.12	0.11*	0.20
	(0.05)	(0.12)	(0.05)	(0.12)
Underweight	−0.70	0.17	0.10	−0.09
	(0.08)	(0.16)	(0.08)	(0.18)
Number of individuals	1,161	1,161	1,161	1,161
Number of siblings	1,042	1,042	1,042	1,042
Chi-square test of difference in OLS and FE is not systematic		11.89		16.50
Probability > chi-square		0.156		0.036

Notes: All models include control variables for 1999–2001 educational attainment, labor market experience, age of youngest child, and age. Race is included in OLS models only. Robust standard errors are reported for Logit. Results for men are insignificant and not reported.
*$p < 0.05$.
**$p < 0.01$.

6. BODY MASS EFFECTS OVER THE LIFE COURSE

Table 6 presents the proportion of younger and older women in healthy, overweight, or obese body mass categories in 2001 by the proportion of women in these categories 15 years earlier. Younger women are between the ages of 25 and 34 in 1986 and older women are between the ages of 35 and 44 in 1986. We compare the proportions between younger and older women to determine if women's movement into and out of healthy weight, overweight, or obesity categories differs by age. Our general conclusion is that it does. About 67% of younger women who were of a healthy weight in 1986 are of a healthy weight 15 years later, whereas only 61% of older women are at a healthy weight 15 years later. Further, of older women who were of a healthy weight in 1986, 9% are obese 15 years later, whereas of younger women who were of a healthy weight in 1986, only 6% are obese 15 years later. Compared with younger women, older women are also less likely to move from

Table 6. Changes in Women's Body Mass by Age: Body Mass Proportions in 1986 by Body Mass Proportions in 1999–2001, by Age (Standard Errors).

	Healthy BMI, 1999–2001	Overweight, 1999–2001	Obese, 1999–2001
Younger women: healthy BMI, 1986	0.67 (0.03)	0.25 (0.03)	0.06 (0.01)
Older women: healthy BMI, 1986	0.61 (0.03)	0.30 (0.03)	0.09 (0.02)
Difference in proportions	0.07**	−0.04**	−0.03**
Younger women: overweight, 1986	0.13 (0.05)	0.41 (0.08)	0.44 (0.08)
Older women: overweight, 1986	0.02 (0.02)	0.50 (0.07)	0.48 (0.07)
Difference in proportions	0.11*	−0.09	−0.04
Younger women: obese, 1986	0.00 (0.00)	0.13 (0.06)	0.88 (0.06)
Older women: obese, 1986	0.04 (0.04)	0.15 (0.07)	0.81 (0.08)
Difference in proportions	−0.04**	−0.03*	0.07*

Note: Younger women are of ages 25–34 in 1986 and older women are of ages 35–44 in 1986.
*$p<0.10$.
**$p<0.05$.

overweight to healthy weight. Clearly, changes in body mass are conditioned by stage in the life course. What is less than clear, however, is if the body mass socioeconomic penalties are conditioned by stage in the life course.

Table 7 presents some evidence that body mass penalties vary across women's stage in the life course. Our age restrictions limit samples sizes and our ability to carry out sibling fixed effects. We rely on a continuous measure of body mass (to reduce problems induced by a small sample size) and we report OLS estimates of the effect of body mass on economic and marital status outcomes. In the analyses reported above, we found little systematic variation between OLS and sibling fixed effects models; nevertheless, we use caution in our interpretation of these OLS models, as unobservable variable bias may differ between older and younger women.

Table 7 reveals that body mass has a stronger negative association with younger women's labor market outcomes. A 1% increase in younger women's body mass leads to a 0.46% reduction in younger women's hourly wages and a 0.67% reduction in women's labor market earnings. The OLS estimates for older women are not significantly different from zero.

Table 7. OLS Models Predicting Women's 1999–2001 Log Wages, Log Earnings, Log Family Income, and Logit Models Predicting Women's 2001 Marital Statuses from a Continous Measure of Women's 1986 Body Mass, by Age Group (Standard Errors).

	Log Wages		Log Earnings		Log Family Income		Married		Divorced	
	Younger	Older	Younger	Older	Younger	Older	Younger	Older	Younger	Older
Body mass	−0.46***	−0.26	−0.67***	−0.22	−0.71***	−1.15**	−0.28*	−0.44**	0.07	0.24
	(0.17)	(0.19)	(0.25)	(0.21)	(0.27)	(0.54)	(0.15)	(0.18)	(0.12)	(0.15)
Number of individuals	263	223	266	225	298	273	298	270	298	270

Notes: Younger women are of ages 25–34 in 1986 and older women are of ages 35–44 in 1986. All models include control variables for 1999–2001 educational attainment, labor market experience, age of youngest child, age, and race. Robust standard errors are reported. Marginal effects are reported for logit marital status models.

*$p < 0.10$.
**$p < 0.05$.
***$p < 0.01$.

Conversely, body mass has a stronger negative effect on older women's marriage market outcomes. For family income, a 1% increase in older women's body mass leads to a 1.15% reduction in family income. And for marriage, a 1% increase in older women's body mass leads to a 0.44% reduction in the probability of marriage. The negative effects of body mass on family income and on the probability of marriage are less strongly negative for younger women.

7. DISCUSSION AND CONCLUSION

Previous research has found that U.S. women pay an obesity wage penalty and men do not (Averett & Korenman, 1996, 1999; Baum & Ford, 2004; Cawley, 2004; Cawley et al., 2005). While all of these studies employ sophisticated techniques aimed at estimating causal effects, all but one of these studies (Cawley et al., 2005) provide estimates that are generalizable only to individuals at the early stages of their lives. Further, studies that have used lagged body mass variables to reduce endogeneity bias have all relied on a 7 year lag. We change the research design of these studies in two respects: we use a much older sample of individuals from the PSID and we employ a body mass measure that is lagged by 15 years. Given these substantial methodological changes, our similarity in findings attests to the robustness of the body mass effect.

Sibling fixed effects models, which we have employed in this study, provide estimates of the body mass effect that are robust to specific forms of unobserved heterogeneity. All unobserved characteristics that siblings share are differenced out of the model, but characteristics that siblings do not share are not differenced out of the model. To the extent that these unobserved, sibling-specific characteristics are correlated with our lagged body mass measure, they could be biasing both traditional OLS estimates as well as our sibling fixed effect estimates. Thus, as with previous research, we can conclude that our estimates approach causality but that they are not the definitive final statement on causality.

Compared with healthy women, obese women pay a 17.51% wage penalty and a 25.06% family income penalty. Assuming that our identification strategy is valid, these are substantial negative causal effects of obesity on women's economic well being. Further, marriage, as one route toward women's economic well being, is less probable for overweight and obese women. Obesity reduces women's probability of marriage by 0.16, compared with healthy weight women, and overweight status reduces women's

probability of marriage by 0.12. As with previous research, we find that men pay no economic or marital status penalties for obesity, with one exception – obese African American men pay a small wage penalty. Finally, we capitalize on the PSID age distribution and find that younger women pay higher labor market penalties for their increased body mass relative to older women, whereas older women pay higher marriage market and family income penalties relative to younger women.

As with any study of women's labor market outcomes, our estimates are subject to a certain amount of selection bias, as women may move into and out of the labor force due to marriage, childbearing, and retirement. Our finding of a stronger obesity penalty for younger women's labor market outcomes, compared with older women's labor market outcomes, may be due to certain older women (namely those with higher socioeconomic statuses and lower body mass indices) selecting out of the labor force and out of our models. On the same token, it is possible that those with higher socioeconomic statuses and lower body mass indices are selecting out of the labor market to bear and rear children when they are younger.

The PSID provides a wealth of data for all stages of the life course (unlike the NLSY which is specifically fielded to one cohort). Nevertheless, disaggregating age and cohort effects within the PSID is difficult. By design, our models control for period effects as we measure body mass and socioeconomic status at standard time periods. Observable age differences, however, could be partially attributable to the fact that our older sample of women came of age more than a decade before our younger sample of women.

A final question worthy of consideration is: how substantial are the effects of increased body mass on women's economic well being? Using estimates reported in Table 3 and sample means on all independent variables, we find that a White obese woman has a predicted wage of about US$ 8.67, whereas a White healthy woman has a predicted wage that is 20% more – US$ 10.48. This difference due to obesity is equivalent to the difference due to almost 2 years of education. The magnitude of the obesity penalty is even greater for family income, as a White obese woman has a predicted value of family income that is 33% less than that of a White healthy woman. This difference in family income due to obesity is equivalent to 2.5 years of education – a substantial difference on any count.

What are the potential causal mechanisms driving significant body mass effects for women? Labor market discrimination may be one causal mechanism that accounts for this relationship. Employers may discriminate against job seekers for both rational and non-rational reasons related to physical appearance. Employers may rationally expect workers who suffer

from obvious health disadvantages to be less productive workers since they may miss work time due to ill health or they may be hampered on the job by a particularly disadvantageous condition. Since health status is not a legally protected category under civil rights law, employers may use some markers of ill health as a proxy for potential productivity – regardless of whether these perceptions are accurate.

Marriage market dynamics may also be predicated on a double standard – heavier women are penalized in the marriage market, as they are less likely to marry than women of a healthier weight, but heavier men are not penalized in the marriage market, as they are equally likely to marry or divorce as men of a healthier weight. These negative effects of body mass on women's likelihood of marriage and divorce become more pronounced as they age. Though our study does not provide definitive evidence as to why women pay these penalties, we do provide evidence of a robust negative effect of increased body mass on women's socioeconomic status. The negative effects of obesity or overweight status persist much longer than previously understood, and they persist across the life course, widening the gender gap in economic well being.

NOTE

1. In addition to body mass, we explored an analysis of height, a different measure of physical appearance. However, all of our estimates using height as a predictor variable were insignificant, and we do not present them.

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