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Abstract

A common problem when using panel data is that an individual's history is incompletely known at the first wave. We show that multiple imputation, the method commonly used for data that are missing due to non-response, may also be used to impute these data that are "missing by design." Our application is to a woman's duration of fulltime employment as a predictor of her risk of first birth. We multiply-impute employment status two years earlier to "incomplete" cases for which employment status is observed only in the most recent year. We then pool these "completed" cases with the "complete" cases to derive regression estimates for the full sample. Relative to not being fulltime employed, having been fulltime-employed for two or more years is a positive and statistically significant predictor of childbearing whereas having just entered fulltime employment is not. The fulltime-employment duration parameter variances are about one third lower in the multiply-imputed sample than in the complete-data sample, and only in the multiply-imputed sample does the employment-duration coefficient attain statistical significance.

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INTRODUCTION

A common problem when using panel data for demographic applications is that the individual's history is unknown or incompletely known at the first wave. In demographic hazard modeling, the length of time an individual has spent in a particular status or activity while exposed to the event of interest is often modeled. State histories may be collected retrospectively at the initial waves of panels surveys to overcome the problem that this duration is otherwise unknown. For example, fertility, marriage, migration, and employment histories are collected at or near the beginning of panel observation in the U.S. Survey of Income and Program Participation (U.S. Census Bureau 2014). State or event histories, however, are not always collected. Even when collected, not all characteristics of the history will be covered. For example, marital and fertility histories alone are not sufficient to identify co-residence histories between two parents and between each parent and child. In a previous treatment of this problem in the U.S. Panel Study of Income Dynamics, Moffitt and Rendall (1995) used a maximum likelihood approach to combine left-censored and non-left-censored spells of single motherhood in separate components of the likelihood. The statistical equivalence of maximum likelihood and multiple-imputation approaches to handling missing data has been noted for cases of data that are missing due to non-response (Schafer and Graham 2002). The ability to separate the imputation step from the analysis step, however, is a major reason to prefer the multiple-imputation approach over the maximum likelihood approach.

Applying further the principles of missing-data analysis, questions that are asked for some individuals and not others in a survey result in a "missing by design" data pattern (Raghunathan and Grizzle 1995). One type of missing-by-design circumstance occurs in combined-survey analysis where the reason that data are missing for some individuals and not

others is because some individuals have been randomly sampled into one survey and whereas others have been randomly sampled into the other survey, and only one of the surveys includes a particular question from which a predictor variable is derived. Multiple imputation has been applied successfully to include such variables in pooled-sample analyses across a diverse set of substantive topics (Gelman, King, and Liu 1998; Rendall et al 2013; Van Hook et al forthcoming).

When the reason that an observation is missing a particular variable is which survey the individual was sampled into, it will usually be clear that the missing at random (MAR) assumption (Little and Rubin 2002), which is necessary for unbiased analysis with multiply-imputed data, will be satisfied. We argue that the MAR assumption will typically be satisfied also when a variable is missing due to left-censoring of a state history when a panel begins. For example, if age 16 is designated as the youngest possible age of employment, we may know for an individual who was age 16 at wave 1 how many years at age 20 he or she has been employed by wave 5 of an annual survey, whereas for an individual who was age 20 at wave 1, this duration is unknown unless an employment history is collected. Whether a particular individual is 16 or 20 at the beginning of the survey may reasonably be treated as random for most analytical purposes, and therefore the unknown employment duration of a 20 year old at wave 1 satisfies the MAR assumption. Despite the analytical attractiveness of their MAR pattern, we are unaware of any previous study that has applied the multiple-imputation approach to left-censored histories in the analysis of panel data.

We use a simple, but nevertheless realistic and widely-applicable example, in which individuals in a panel survey contribute either one or two waves of employment history with which to predict a partnered first birth. Our data source is the Poland country survey of the

European Union Survey on Income and Living Conditions (EU-SILC, Eurostat 2011a). Poland follows the standard format of including rotating panels, each including four annual waves. Using all four waves, we can use employment status in the first and second waves to predict births reported between the third and fourth waves. Using multiple imputation, however, we are additionally able to include birth-exposure intervals between the second and third wave in a model that uses employment status from *two* previous years. Compared with a model estimated with only birth-exposure intervals between the third and fourth waves, our parameter estimates are then derived from a sample whose number of person-year observations is twice as large. We discuss conditions under which the standard errors around the coefficients for duration employed (1 year only versus 2 or more years) may be accordingly reduced and find empirically that they are indeed substantially lower. Substantive implications from the multiply-imputed model results include that, whereas a woman's having been employed fulltime for at least the last two years is statistically associated with a higher propensity to give birth, her having just entered a spell of fulltime employment is not. This result holds, moreover, when controlling for her partner's employment status, a variable that is available in our panel data but that is typically not available in retrospective data.

DATA AND METHOD

The European Union Survey on Income and Living Conditions (EU-SILC) is a set of more than 30 country surveys conforming to a common format and content specified by Eurostat (Eurostat 2011a). The EU-SILC was created in 2003 as a replacement for the European Community Household Panel (ECHP). It gathers harmonized and comparable data at the individual and at the household level on income and living conditions as well as on many individuals' demographic

and socio-economic characteristics (sex, age, education, labor market position, etc.). The EU-SILC is composed of two datasets – one cross-sectional and one longitudinal (Eurostat 2011b). Subgroups of individuals observed in the cross-sectional dataset are followed up over several years to form the longitudinal dataset. The longitudinal design consists of a rotational panel in which individuals are observed annually for a standard maximum period of four years. The EU-SILC covers the majority of European countries for the years 2003 to 2012, although some began later than 2003, notably in 2005 when 10 new countries joined the European Union. For the present study, we use the longitudinal EU-SILC data for one country only, Poland, whose survey began in 2005. We choose Poland due to its relatively large number of observations (the number of observations available for each country reflects its population size, which was 38.5 million in 2011, Central Statistical Office of Poland 2012), and because the quality of the data has been assessed favorably (Iacovou, Kaminska, and Levy 2012). In particular, its wave-to-wave retention rate of approximately 90% is among the highest among EU-SILC countries. This retention rate exceeds that of all other larger countries in the EU-SILC. Selection into the Polish sample occurs annually, with each new sample followed for a further three waves. A stratified random sampling design ensures that sufficient sample sizes are attained for EU-designated “NUTS2” regions that are classified by degree of urbanization (Eurostat 2011a, p.33). We use sample weights that adjust for this design in the univariate descriptive statistics of our study.

We limit our analyses to partnered Polish women aged 18 to 39 who are observed for three or four consecutive waves between the years 2005 and 2012 and who are childless when entering the survey. The upper age restriction is needed so that we may reasonably approximate a woman’s being of parity 0 by her having no co-resident children. The restriction to individuals observed for a minimum of three waves is necessary since interviews usually take place during

the first half of each year children born in the third and the fourth quarters of each year are generally reported at the interview of the year after the birth. Two consecutive years of interviews are then needed to identify all births that occur in one calendar year, and at least one more preceding wave is needed to observe the mother's (and her partner's) labor market status before exposure to conception and fertility. For those individuals who are observed for three waves, the latter two waves are designated t and $t + 1$ and serve to detect a first child arrival in the calendar year of wave t . Wave $t - 1$ is used to observe the woman's employment status and that of her partner. For those individuals who are observed for four waves and for whom no birth occurs in any of the calendar of the years of the first two waves, the woman's employment status before birth exposure is observed in both waves $t - 2$ and $t - 1$.

An alternative data source for analyzing fertility by employment status in Poland, not used in the present study, is the retrospective 2006 Employment, Family, and Education Survey (EFES). The EFES was used by Matysiak (2009), who noted a disadvantage of that survey is its high rate of non-response among women's partners (more than 50%), leading the author to analyze only women's characteristics as predictors of their fertility. Contemporaneously-collected information on partners is a major advantage of panel data for the analysis of fertility by employment status and earnings (e.g., Adsera 2011).

Multiple Imputation (MI) for observations that are incomplete "by design"

We first describe the general "missing by design" setup and results already known with respect to efficiency gains from multiple imputation of missing data, and then proceed to the special case of incomplete information on employment-spell durations as predictor variables. We first specify a model with outcome variable Y_t for a birth in calendar year t as a function of predictor

variables X_{t-2} and X_{t-1} observed respectively at times $t - 2$ and $t - 1$. The birth event is the result of a *conception* that will in most cases have occurred between $t - 1$ and t . Our model therefore avoids predictor variables observed at time t , since it may not be clear what was the causal ordering between a conception carried to term, Y_t , and a woman's characteristics at time t , X_t . We use the binary logit model, $LOGIT[p] = \ln[p / (1 - p)]$, for the regression:

$$LOGIT[\Pr\{Y_t|X_{t-2}, X_{t-1}\}] = \beta_0 + \beta_1 X_{t-2} + \beta_2 X_{t-1} \quad (1)$$

The data available to us to estimate this model are “complete” observations $\{Y_t, X_{t-2}, X_{t-1}\}_{i=1}^{N_1}$ and “incomplete” observations $\{Y_t, X_{t-1}\}_{j=1}^{N_2}$. We assume that the substantive model represented by equation (1) applies equally to the complete and incomplete observations. This allows us to apply standard multiple imputation procedures to combine them in the analysis. Our specific application is as follows. We allow woman's employment status E to have an effect on Y_t based on its values at both times $t - 2$ and $t - 1$, E_{t-2} and E_{t-1} . The other predictor variables in vector X , which we will henceforth denote by Z , all have effects on Y_t only from their values at time $t - 1$. Our variant of equation (1) is therefore:

$$LOGIT[\Pr\{Y_t|E_{t-2}, E_{t-1}, Z_{t-1}\}] = \beta_0 + \beta_1 E_{t-2} + \beta_2 E_{t-1} + \beta_3 Z_{t-1} \quad (1a)$$

We ignore item non-response, which is anyway very low for our variables of interest in the Poland EU-SILC, and assume that we have only “complete” observations $\{Y_t, E_{t-2}, E_{t-1}, Z_{t-1}\}_{i=1}^{N_1}$ and “incomplete” observations $\{Y_t, E_{t-1}, Z_{t-1}\}_{j=1}^{N_2}$. As a consequence, the pattern of missingness is monotone (Little 1992) and we are able to apply a chained-equation multiple imputation approach (Raghunathan et al 2001) that allows for the imputation of binary, count, or continuous data. We use the set of complete observations to first estimate the imputation equation:

$$LOGIT[\Pr\{E_{t-2}|E_{t-1}, Z_{t-1}, Y_t\}] = \gamma_0 + \gamma_1 E_{t-1} + \gamma_2 Z_{t-1} + \gamma_3 Y_t \quad (2)$$

We then apply random draws from the posterior distribution of parameter estimates $\hat{\gamma}_0, \hat{\gamma}_1, \hat{\gamma}_2, \hat{\gamma}_3$ to the incomplete data $\{E_{t-1}, Z_{t-1}, Y_t\}_{j=1}^{N_2}$ to derive an arbitrarily large m values of E_{t-2} (we set $m = 20$) to produce *completed* data $\{\{Y_t(k), E_{t-2}(k), E_{t-1}(k), Z_{t-1}(k)\}_{j=1}^{N_2}\}_{k=1}^m$. We then concatenate the *complete* data $\{Y_t, E_{t-2}, E_{t-1}, Z_{t-1}\}_{i=1}^{N_1}$ to each instance of *completed* data and estimate the analysis equation (1a) m times. These m estimates of (1a) are combined using the standard multiple-imputation algorithms (Little and Rubin 2002) to produce a set of parameters with standard errors $\{\hat{\beta}_0, \hat{\beta}_1, \hat{\beta}_2, \hat{\beta}_3; SE(\hat{\beta}_0), SE(\hat{\beta}_1), SE(\hat{\beta}_2), SE(\hat{\beta}_3)\}$ that adjust for the uncertainty introduced by imputation of E_{t-2} to the incomplete person-year observations. This sequence of procedures may be performed with standard package software. We use SAS PROC MI and PROC MIANALYZE (SAS Institute 2008a, 2008b).

It is helpful to know generally what may be the efficiency gains from combining the complete and incomplete data. Equation (1) has been used in prior work to derive or estimate these expected efficiency gains. Little (1992) derived results in the linear regression case and White and Carlin (2010) did so additionally for the logistic regression case. A key parameter in evaluating efficiency gains obtained through multiply imputing is the “fraction missing.” In equation (1), there are only two predictors, one observed in both the complete and incomplete data, X_{t-1} , and another, X_{t-2} , observed only in the complete data. Following the terminology of White and Carlin, we define the fraction with missing values of X_{t-2} by $= N_2/(N_1 + N_2)$. The increase in sample size when both the complete and incomplete data are used in the estimation results in efficiency gains (reductions in standard errors or variances about the parameters) over an analysis that would estimate equation (1) from the complete data only. The larger the fraction

missing, the larger will be the efficiency gains. The largest variance reductions will be for $Var(\hat{\beta}_2)$, the parameter for which the regressor variable (X_{t-1}) is observed in both the complete and incomplete data. Variance reduction about $Var(\hat{\beta}_2)$ depends not only on the fraction missing π , however, but also on the correlation ρ_{12}^2 between X_{t-2} and X_{t-1} and on the partial correlation of Y_t and X_{t-2} given X_{t-1} , $\rho_{1y.2}^2$. Analytical expressions for the variance reductions about $Var(\hat{\beta}_2)$ and $Var(\hat{\beta}_1)$ are available in the linear case, originally from Little (1992). The expression for the proportion by which $Var(\hat{\beta}_2)$ is reduced by adding left-censored observations in the linear regression case is given by (White and Carlin 2010, p.2922):

$$\pi(1 - \rho_{1y.2}^2)[1 - \rho_{12}^2(1 - 2\rho_{1y.2}^2)] \quad (3)$$

In the special case of no correlation between X_{t-2} and X_{t-1} and when X_{t-2} has no association with Y_t independent of variation in X_{t-2} , then $Var(\hat{\beta}_2)$ reduces by the maximum amount, equal to the fraction of left-censored observations, π . In this case, however, we could estimate β_2 without the need for MI. Instead we would simply pool the complete and incomplete data and estimate $LOGIT[\Pr\{Y_t|X_{t-1}\} = \beta_0 + \beta_2 X_{t-1}]$. In the more general case of $0 < \rho_{1y.2}^2 < 1$ and $0 < \rho_{12}^2 < 1$, the theoretical result is that MI will always result in a reduction in $Var(\hat{\beta}_2)$ in the linear regression case because both $(1 - \rho_{1y.2}^2)$ and $[1 - \rho_{12}^2(1 - 2\rho_{1y.2}^2)]$ will always be less than 1. Because they are greater than zero, the theoretical reduction in variance will be smaller than the fraction of left-censored observations, π . White and Carlin conduct simulations and real-data analyses and find that the variance reductions are typically close to (i.e., not much less than, and in some cases greater than) the fraction of missing values in both linear and logistic regression cases.

Reductions in $Var(\hat{\beta}_1)$ are of particular interest because the coefficient $\hat{\beta}_1$ is associated with the variable X_{t-2} for which data are partly missing. The reductions in variance about $\hat{\beta}_1$ are not generated directly, because only in the complete data is there any information about the relationship between Y_t and X_{t-2} . Instead, the only reduction in the variance of $Var(\hat{\beta}_1)$ arises from the partial correlation of Y_t and X_{t-2} given X_{t-1} , $\rho_{1y.2}^2$. That is, efficiency may be gained by adding observations from the incomplete data (in our case, the left-censored observations) due to their indirectly providing information about the relationship between Y_t and X_{t-2} . This is because X_{t-2} and X_{t-1} are correlated and this correlation carries with it information about variation in Y_t through providing information about the relationship between Y_t and X_{t-1} . The analytic expression for the variance reduction in $Var(\hat{\beta}_1)$ in the linear regression case is, again originally from Little (1992):

$$2\pi\rho_{1y.2}^2(1 - \rho_{1y.2}^2) \quad (4)$$

In most cases, the value of this partial correlation will be low, and therefore negligible reduction in of $Var(\hat{\beta}_1)$ is expected (e.g., White and Carlin 2010). In our case, however, the correlations between X_{t-2} and X_{t-1} and the partial correlation of Y_t and X_{t-2} given X_{t-1} are expected to be high, and therefore the reduction in $Var(\hat{\beta}_1)$ potentially substantial. This is because our X_{t-2} and X_{t-1} are composites of employment status in times $t - 2$ and $t - 1$.

In particular, we consider three durations of fulltime-employment spells in progress at the time of birth exposure D_l : 0, 1, and 2 or more years. The reference category is $D_0 \equiv \{1 \text{ if } E_{t-1} = 0 \text{ and } 0 \text{ if } E_{t-1} = 1\}$ and therefore requires only information from $t - 1$. To code the alternate categories, of duration exactly 1 year, D_1 , and duration of two or more years, D_2 , information at both times $t - 2$ and $t - 1$ is required, since $D_1 \equiv \{1 \text{ if } E_{t-1} = 1 \text{ and } E_{t-2} = 0\}$ and $D_2 \equiv \{1 \text{ if } E_{t-1} = 1 \text{ and } E_{t-2} = 1\}$. The analysis model we estimate is:

$$\text{LOGIT}[\text{Pr}\{Y_t|D_2, D_1, Z_{t-1}\}] = \beta_0 + \beta_1 D_2 + \beta_2 D_1 + \beta_3 Z_{t-1} \quad (1b)$$

Since reference duration category D_0 and alternate duration categories D_1 and D_2 use both complete and multiply-imputed (“completed”) data for their coding, the expected efficiency gains in the estimation of β_1 and β_2 are *a priori* unknown. They are expected to be less, however, than the efficiency gains in the estimation of β_3 . This is because Z_{t-1} is assumed to be non-missing for all observations, whereas both D_1 and D_2 are constructed partially from observed data and partially from multiply-imputed data.

RESULTS

The sample consists of person-years exposed to a first birth among partnered, parity-0 women (see Table 1). Women’s mean age of exposure to first birth was 28.4 (from an observed age range of 20 to 39). We code employment status as either fulltime-employed or not for both the woman and her partner: 74.5% of women were fulltime-employed the year immediately preceding a person-year of exposure to first birth, and 70.8% were fulltime-employed two years before the person-year of exposure to first birth; 85.9% of their partners were fulltime-employed the year immediately preceding their calendar year of exposure to first birth. Note that for only 200 person-years do we observe the woman’s employment status two years before their calendar year of exposure to first birth. Of a total of 671 person-year observations (maximum of two per woman), 323 were full-time employed in the year before birth exposure ($t - 1$) and did not have employment status observed in the year before that ($t - 2$) because they had not yet entered the panel. These are the left-censored spells. The “complete” data consist of 348 person-years, and thus the fraction with left-censored spells, referred to more generally as the fraction missing, is 0.481.

Regression results are presented in Table 2. We first use conventional methods to estimate models predicting a birth. These either simplify the specification of the employment-status predictor to use status only at $t - 1$ (Model 1) or use the fuller specification of employment status at both $t - 1$ and $t - 2$, but necessarily for a person-year sample only half as large (Model 2). In the Model 1 estimation, in which all 671 person-years are used, but for which the specification of employment status is reduced to one prior year, being fulltime-employed at $t - 1$ is a statistically-significant predictor of giving birth, associated with a 1.92 greater odds of giving birth compared to having not been fulltime-employed in the prior year. This result is consistent with Matysiak's (2009) finding using retrospective data in the 2006 EFES survey, in which she also used employment status only in the year immediately before exposure.

When in Model 2 we distinguish between 1 year only of fulltime employment and 2 or more years of fulltime employment, we find that having been fulltime-employed 2 or more years ("duration 2+") is associated with a 1.90 greater odds of giving birth compared to not having been fulltime-employed in the prior year ("duration 0"). This is statistically significant, however, only at the 0.10 level ($p = 0.07$). Women who entered fulltime employment in the most recent year ("duration 1") did not have a statistically-significantly different likelihood of giving birth compared to women who were not fulltime-employed in the most recent year. These results are suggestive of duration of fulltime employment being a critical factor for predicting a partnered woman's first birth. Using these conventional estimation methods, however, we are only able to include employment duration in the model at the cost of eliminating almost half of an already relatively small sample size, thereby rendering both employment-duration coefficients non-significant at conventional thresholds.

Our preferred model is Model 3, in which all 671 person-year observations are used, and with a specification of fulltime employment that distinguishes 0, 1, and 2+ years' duration. This is the model made possible by multiply-imputing values of the fulltime-employed variable for the 323 person-years in which the woman was observed to be fulltime-employed at time $t - 1$ and was not observed at time $t - 2$. Our coefficient values are reassuring similar to those of Model 2, which used only complete data, indicating unbiased estimation with the multiply-imputed data. Model 3, however, has substantially lower standard errors than Model 2. This includes the standard error for the coefficient of greatest substantive interest to us, fulltime-employed for 2 or more years. Its standard error reduces from 0.36 in Model 2 to 0.29 in Model 3. Being fulltime-employed for 2 or more years is associated with a 2.14 greater odds of giving birth compared to not having been fulltime-employed in the prior year ("duration 0"), and the coefficient is now significant at the 0.01 level ($p = 0.009$). Entering fulltime employment in the most recent year ("duration 1") was again not statistically-significant. In none of the three models is the coefficient for partner's fulltime employment statistically significant, although it is in each case positive and with estimated magnitudes similar to those for the woman's fulltime employment.

From prior theoretical and simulation studies cited above, the standard errors for the variables for which there are no missing values were expected to be much lower in the Model 3 than in the Model 2 estimates. In the linear regression case, the maximum theoretical reduction in variances (that is, the square of the standard error) is the proportion of cases that are missing, which in our case is 0.481 (being the fraction of person-years for which we cannot define the fulltime-employed duration as 0, 1, or 2+ years, see again Table 1). Although there is no such simple analytical expression in the logistic regression case, we see that the reduction in variances

for the age and age squared coefficients are 0.508 and 0.521, and so quite close to the fraction missing. The proportionate reduction in variance about the partner's employment status is 0.457, again close to the fraction missing. Of greatest interest is the reduction in variances about the two employment spell duration coefficients. We expected that the reduction would be less than for those coefficients for variables that are constructed only from information at time $t - 1$, and therefore non-missing, but we did not know by how much, nor even if there would be any substantial reduction in variances over the complete-data estimation. The proportionate reductions in variances are respectively 0.325 and 0.291 for the coefficients for fulltime employed two or more years and for fulltime employed only one year. These reductions are substantially less than the fraction missing, as expected, but are nevertheless surprisingly large magnitudes.

CONCLUSION

The use of multiple imputation in demographic research now spans at least 20 years (for early uses, see Freedman and Wolf 1995; Goldscheider et al 1999; and Sassler and McNally 2003). The standard use that is made of multiple imputation is to correct for various forms of non-response (see also Johnson and Young 2011). In the present study, we have argued that a promising additional use for multiple imputation is for data that are instead "missing by design" (Raghunathan and Grizzle 1995), in which the survey's design means that only a fraction of the sample is asked a particular question. A major advantage of using MI for data that are missing by design is the expected validity of the assumption that data are missing at random (MAR). This assumption is both a requirement of, and is the biggest challenge for, the successful (unbiased) implementation of the method of multiple imputation (Schafer and Graham 2002). MI for

missing-by-design data patterns has previously been successfully implemented in combined-survey analysis (Gelman, King, and Liu 1998; Rendall et al 2013; Van Hook et al forthcoming), where the reason that data are missing for some individuals and not others is because some individuals have been randomly sampled into one survey and others have been randomly sampled into the other survey. We argue that left-censored spells beginning in the first wave of panel data are similarly likely to meet this MAR requirement because missingness again results from the random sampling process. The value of the MI method as applied in the present study is further enhanced by the ubiquity of left-censored spells beginning in the first wave of panel data. In short panels and in panels that sample from populations rather than from cohorts, left-censored spells of employment and family status, for example, occur for almost every adult observed in the panel.

In the present study's application of the MI method, we examined the gains that may be realized by multiply imputing a single additional year of employment status before the first wave of the panel. This amount of imputation was the maximum possible given the very short panel of our example data, the four-wave EU-SILC. Nevertheless, doing so allowed us to conduct a simple test of the hypothesis that women are more likely to begin childbearing after first obtaining *stable* employment. Using conventional methods to conduct this test would have meant using only half the number of person-year observations that we were able to use in our multiply-imputed data analysis. Alternative data or analytical methods that could instead have been used all have limitations. To proxy for stability of employment as a predictor of fertility when no employment history was collected in the European Community Household Panel (ECHP), Santarelli (2011, p.322) used a variable for type of employment contract (permanent or temporary), but this type of contract is likely to be country-specific in its meaning and a variable

for it is not consistently available across countries. Özcan, Mayer, and Luedicke (2010) criticized studies that ignore problems of left-censoring and instead used a cross-sectional survey, the German Life History Survey, with employment and demographic histories. The life histories collected in cross-sectional surveys, however, have major limitations, notably that they often include little or no recording of the employment statuses of current and former partners.

In the empirical part of our study, we found that a woman's being newly fulltime-employed was not predictive of a first birth whereas being fulltime-employed for two or more years was strongly predictive of a first birth. In terms of the estimated coefficient sign and magnitude, we found this result to be consistent between estimation using the larger set of person-years that we "completed" by multiple imputation and estimation using the smaller set of "complete" person-years in which the employment spell of up to two years was observed. This concordance in estimates between the complete-data and multiply-imputed data analyses is reassuring from the perspective of protecting against any bias that might be induced by the multiple imputation of data from the complete observations to the incomplete observations. Only in the analysis with the multiply-imputed data, however, was this key substantive coefficient statistically significant at conventional levels ($p < .05$). The magnitude of variance reduction about the parameter was around one third. To have obtained a variance reduction of this size represents a substantial payoff to having multiply-imputed a variable for employment status two years before the birth exposure for such a large fraction of our person-year sample. This large amount of variance reduction, moreover, was not easily knowable in advance of our conducting the study, since the left-censored spell case of the present study is qualitatively different from those for which previous work has derived expected variance reductions (e.g., White and Carlin 2010). We attribute the large variance reduction to the fact that for every observation at least

some information was available on the length of the spell. Future work, however, might profitably investigate the different amounts of variance reduction that may be realized under different types and magnitudes of missing versus non-missing information in left-censored histories.

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Table 1 Descriptive statistics and numbers of observations, partnered parity-0 Polish women, 2005-2012

Descriptive statistics (person-years, weighted)

	Mean	Standard deviation	Sample size
woman's age	28.7	4.4	671
woman's fulltime employment in t-1 (proportion)	0.745	0.436	671
woman's fulltime employment in t-2 (proportion)	0.708	0.456	200
partner's fulltime employment in t-1 (proportion)	0.859	0.349	671

Numbers of observations (unweighted person-years)

Left censored observations: fulltime employed at t-1, not observed at t-2	323
Non-left-censored observations	348
All observations	671

Fraction missing full-time employment duration (left-censored)	0.481
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Note: all observations have valid values of age and partner's employment status at t-1 and of birth between t and t+1

Source: European Union Survey of Income and Living Conditions, Poland 2005-2012

Table 2 Logistic regressions of birth in year t, before and after imputing fulltime employment status in t-2, partnered parity-0 Polish women, 2005-2012

	Model 1				Model 2				Model 3				Reduction in variance, Model 2 to Model 3*
	complete data, reduced specification				complete data, full specification				complete and multiply- imputed data, full specification				
	Stand- Odds		ard		Stand- Odds		ard		Stand- Odds		ard		
	Estimate	Ratio	Error	p-value	Estimate	Ratio	Error	p-value	Estimate	Ratio	Error	p-value	
Intercept	-0.056		4.02	0.989	-2.796		5.85	0.633	0.536		4.16	0.898	0.494
fulltime-employed at t-1	0.652	1.92	0.26	0.013									
fulltime employed at t-1 and t-2 ("duration 2+ years")					0.641	1.90	0.36	0.071	0.763	2.14	0.29	0.009	0.325
fulltime employed at t-1 but not t-2 ("duration 1 year")					-0.155	0.86	0.67	0.817	0.209	1.23	0.57	0.713	0.291
age	-0.069		0.29	0.810	0.135		0.42	0.748	-0.105		0.30	0.721	0.508
age squared	-0.001		0.005	0.906	-0.004		0.007	0.557	0.000		0.005	0.986	0.521
partner fulltime employed at t-1	0.550	1.73	0.34	0.111	0.549	1.73	0.47	0.246	0.566	1.76	0.35	0.104	0.457
sample n	671				348				671				

Notes:

* calculated by squaring the Standard Errors and taking the proportionate reduction in these variances about the parameter estimate from Model 2 to Model 3

Source: European Union Survey of Income and Living Conditions, Poland 2005-2012