Employment Impacts on Entry to Parenthood in Different Family-Policy Regimes

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ABSTRACT

Our goal was to explore how women and men's employment interacts with family policy and social norms to produce differences in gender inequalities in the relationship of employment to first birth. Using comparable panel data from the 2000s across six high-income countries, we estimated identical models of individual employment on partnered women's and men's entry to first parenthood, and on unpartnered women's entry to first parenthood. Two countries each were from 'dual-earner' (Norway and France), 'liberal' (Australia and the United States), and 'conservative' (Germany and Switzerland) family-policy regimes. We tested three hypotheses generated from theory of reproductive polarization, in which family policy is claimed to play a central role in generating or mitigating socio-economic heterogeneity in family formation. We found support overall for our hypotheses. Women and men in 'dual-earner' regimes, in particular, had higher rates of entry to first parenthood in the year prior to fertility exposure when 'full-year, full-time' employed compared to those employed little or not at all in the year prior to fertility exposure. We found substantial variation between 'liberal' and 'conservative' regimes in fertility responses to employment, with unexpectedly positive relationships of being 'full-year, full-time' employed to first birth rates among German women, in contrast to expected negative relationships of employment to first birth rates among Australian women, especially when unpartnered. Partnered women's proportions in full-time, full-year employment were surprisingly as much as 15 to 25 percentage points lower than partnered men's proportions across the five countries for which we made this comparison. This, and an exceptionally low proportion of women who were partnered in Germany, led us to conclude that modeling of the selection into

partnership and employment may provide valuable complementary understanding of the relationship of employment to first parenthood across family-policy regime types.

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INTRODUCTION

Studies of socio-economic status (SES) differentials in family formation in high-income countries point to what has been remarked in the case of the U.S. as "…a troubling divergence in the family patterns of Americans according to education and income…" (Cherlin 2010, pp.403-404). This conclusion follows a previous review of SES disparities in family formation (McLanahan 2004) that drew contrasts between two broad groups: a disadvantaged group characterized by earlier and often non-marital family formation and an advantaged group characterized by later family formation with childbearing occurring within marriage. This polarized characterization of U.S. fertility is supported by recent empirical findings (Shang and Weinberg 2013). Similarly-steep SES gradients in family formation as in the U.S. have been documented in other countries. These include a much larger reduction in marriage among less educated than among more educated women in Australia and New Zealand (Heard 2011), and findings of more cohabitational and less marital childbearing among less educated women in a study of several European countries (Perelli-Harris et al 2010).

There is also, however, evidence indicating that the phenomenon of polarization of family formation along SES lines is not a universal one. Perelli-Harris et al's (2010) study noted Sweden to be an exception in exhibiting little difference in cohabitational versus marital fertility by SES. This is consistent with smaller SES differences in fertility across Nordic countries (Rønsen and Skrede 2008). Regarding overall fertility, Toulemon, Pailhé, and Rossier (2008) describe a *narrowing* of women's and men's differences by educational attainment in France. One explanation that has been offered (Rendall et al 2009) is that "family-friendly" policy regimes such as those of France and the Nordic countries promote more homogeneous ages of entry to parenthood by SES than do 'liberal' regimes (Esping-Andersen 1999) such as those of the United States and the other high-income English-speaking countries. Rendall et al (2009) found that while the timing of first births became more strongly differentiated by women's occupational levels over time in Great Britain, it did not do so in France.

Although theoretical treatments of family-policy regimes have emphasized the contrast between those that do versus do not facilitate the combining of employment and parenthood, there are two distinct alternatives to the "family-friendly" regimes of the Nordic countries and France. These alternatives correspond to the 'liberal' and 'conservative' regimes of Esping-Andersen's (1990; 1999) seminal work differentiating between broad welfare policy types. Morgan (2003) and Joshi (2002), among others, note that overall fertility rates are much higher in 'liberal' regimes than in 'conservative' regimes. They describe the ways that employment and parenthood are reconciled in 'liberal' regimes as including a diverse mix of part-time work, spells out of the workforce with re-entry in a different job, and the use of private-sector childcare. The means-testing structure of family policy in 'liberal' regimes such as the U.S. also allows for an alternative, non-marital fertility path among women whose low educational attainment gives them poor labor-market prospects (Rosenzweig 1999; Hoffman and Foster 2000).

The goal of the present study is to compare the effects of women's and men's own employment on first parenthood across the above three regime types. Our overall prediction is that in 'dual-earner' regimes, being full-time, full-year employed in the previous year will be associated with higher probability of a first birth among women, and that the relationship of employment to having a first partnered child will be similar between men and women; in contrast, we predict an inverse effect of employment on first birth for women in 'liberal' and in 'conservative' countries. We further predict that employment will be positively related to having a first partnered child among men in all countries.

To anticipate our findings, our predictions are mostly supported, but with some unexpected findings and some non-significant findings. Norwegian and French findings were largely in line with predictions, with little or no work in the previous year being associated with a lower likelihood of first birth among both partnered women and (in Norway) partnered men. Australian findings were somewhat in line with predictions, with part-time or part-year work in the previous year being associated with a higher likelihood of first birth among partnered women

than was full-time, full-year work, and little or no work in the previous year being associated with the highest likelihood of first birth among unpartnered women. Unexpectedly, in Australia part-time or part-year work in the previous year was associated also with a higher likelihood of first entry to parenthood also among partnered *men* than was full-time, full-year work, although this was only marginally statistically significant. Also unexpectedly, in Germany full-time, full-year work in the previous year was associated with a higher likelihood of first birth among both partnered and unpartnered women. Evidence of very high unpartnered proportions of 18-34 year old women in Germany, however, indicates a possible response of delaying or foregoing entry to marriage or cohabitation as a means of delaying or foregoing entry to motherhood in an environment unfavorable for combining employment and motherhood. The U.S. and Switzerland were mostly remarkable for their lack of statistically-significant coefficients for employment, in part due to their smaller sample sizes.

LITERATURE REVIEW

Family policies that facilitate the combining of employment and motherhood through the provision of subsidized child-care and maternity-leave compensation and rights to return to work after the leave are believed to be especially important for mitigating role incompatibility (Pampel 2001; Castles 2003). However, whereas strong family-policy effects on employment continuity have been found (Ruhm 1998; Stier et al. 2001; Rønsen and Sundström 2002; Thévenon and Solaz 2013), the impacts of family policy on fertility have found to be mostly relatively small or short-lived (van de Kaa 2006, Gauthier 2007, but see also McDonald 2006 and Toulemon 2011). In a recent study that directly compared labor-force-participation and fertility effects of family policies, the former effects were found to be substantially larger (Del Boca et al 2009).

The effects of (family-friendly) family policies on reducing fertility *differentials*, including by SES, may be greater than any effects they have on fertility levels. In two previous cross-national analyses, Rendall and colleagues (Rendall et al 2009; 2010) found evidence for

potentially large impacts of family-policy regime type on *differences by SES* in family formation. They found growing differences by education and occupation in the timing of first birth and in childlessness in countries (including the U.S. and the U.K.) with family-unfriendly policy regimes but not in countries with family-friendly regimes. Their findings are consistent with theoretical and empirical work indicating strong incentives for women in family-friendly regimes to first secure regular employment before beginning childbearing (Gustafsson 2001). Gustafsson and colleagues (Gustafsson et al. 1996; Gustafsson and Wetzels 2000) argue, with supporting evidence from cross-national analyses of example countries in different family-policy regime types, that a well-integrated set of family-policy provisions provides strong incentives for women across socio-economic levels to first secure permanent employment before entering parenthood, and then to proceed quickly to childbearing with their jobs held for their return after each maternity leave. This is expected to lead to more homogeneous ages of first birth by education and income by allowing lower-income women to afford institutional childcare. This prediction is consistent with theoretical treatment by Ermisch (1989) and with empirical and theoretical treatment of childcare costs among overall costs of children by DiPrete et al (2003). It has also been suggested, however, that the use of childcare subsidies may also serve only to change the cutoff point for which women can afford to use paid childcare, still leaving the lowest earners unable to take advantage of these provisions (Thévenon and Gauthier 2011).

Empirical studies examining the potential effects of family-policy on reducing or increasing socio-economic differentials in fertility are relatively few. Meron and Widmer (2002) found postponement of childbearing in response to unemployment in France to be at least as great among low-educated women as among women with higher education levels, consistent with the arguments of Gustaffson and colleagues that women at all SES levels will delay parenthood until obtaining secure employment in policy environments that universally subsidize and support the combining of employment and motherhood. Aasve, Billari, and Spéder (2006) found that Hungary's switching from a universalistic to means-testing childcare benefits policy,

and then back again to a universalistic policy, had large effects only on the childbearing of more educated women, with the temporary switch to a means-testing regime resulting in a widening of socio-economic differentials in childbearing. Additional indications of the potential power of institutional context to influence fertility distributions are seen in findings of less marked socio-economic differentials in higher-order births in the universalistic countries of Scandanavia and France than elsewhere in Europe (Ekert-Jaffé et al 2002; Callens and Croux 2005).

The focus of studies of the effects of 'conservative' family policy has been on delayed and reduced fertility in these regimes (e.g., Kohler, Billari, and Ortega 2002; Sánchez-Barricarte and Fernández-Carro 2007), and not on potential SES differentials in these patterns of delay and decrease. Even when the word "polarization" has been used (e.g., Dorbritz 2008), it has been used in a purely demographic sense in which women choose to either have no children or two or more children under "family-unfriendly" regimes from the 'conservative' group, such as those of Germany and Switzerland. Nevertheless, Schulze and Tyrell (2002) have characterized Germany's high rates of childlessness as being socio-economically constrained by "family-unfriendly" policies that make it much more costly for higher- than lower-educated women to become mothers, a phenomenon they described as "reproductive polarization".

Following up on this characterization in a study of fertility timing by education including three 'conservative' countries in Southern Europe, Rendall et al (2010) found that the most extreme patterns of increasing delay in first birth across countries occurred among high-educated women in these 'conservative' countries. They found at the same time, however, persistence in patterns of early first births among groups of low-educated women in these countries. The latter patterns were in parallel with patterns of persistence in early first births in the 'liberal' countries of Great Britain and the United States, and were in contrast to declines in early first births among low-educated women in France and Norway. Gonzalez and Jurado-Guerrero (2006) also found contrasts between France and three conservative countries, Germany, Italy, and Spain, that pointed to likely family-policy regime effects. They summarized

(p.339) the contrast between France and the three ‘conservative’ countries as being consistent with a dual-earner versus a male-breadwinner model: “…the peculiarity of the French [context] seems to be that being in paid employment favours motherhood more than being a housewife.”

Most studies of SES differentials in family-demographic events have used educational attainment as a proxy for SES (e.g., Rendall et al 2010; Heard 2011; Perelli-Harris et al 2010, 2012), and relatively few have examined employment and the labor market as a mechanism through which educational attainment may generate these differentials (for notable exceptions, see Adsera 2004, 2005). The effects of employment status, in general, on family formation have been found to vary substantially across countries (Sobotka, Skirbekk, and Philipov 2011). The effects of *women’s* employment status are especially variable across countries. This is argued to be at least partly due to differences in the degree to which a country’s family-policy provisions facilitate the combining of motherhood and employment (Matysiak and Vignoli 2008). Increases in employment and earnings *uncertainty* are also expected to affect family formation, and this too seems to vary by family-policy context. Schmitt (2012) found greater delays in becoming a mother among employed women in the U.K. than in Germany. Pailhé and Solaz (2012) found that employment uncertainty reduced fertility in Germany whereas it delayed but did not reduce fertility in France. Those authors speculated that this difference in findings across countries may be due to the mitigating effects of France’s strong family and welfare policy provisions.

We know of no previous study that compares women’s and men’s entry to parenthood by employment across the three main family-policy regime types. By doing so here for the decade of the 2000s, the present study addresses the overarching research question of whether the relationship between employment and family formation differs across family-policy and macro-social contexts in ways predicted by theory of “reproductive polarization”. We frame our analysis in terms of differences by gender in the effects of employment on first parenthood, and differentiate women’s entry to parenthood additionally by whether they were partnered or

unpartnered in the year before exposure. We describe our specific predictions in hypothesis form following presentation of our data and estimation approach.

DATA AND METHOD

Our research question is addressed in the present study with estimates from panel data of the relationship of employment to first parenthood, and how this differs between men and women, across six countries. These consist of two countries from each of three family-policy regime types: France and Norway, representing the 'dual-earner' regime type; Australia and the United States, representing the 'liberal' regime type; and Germany and Switzerland, representing the 'conservative' regime type. Although Esping-Andersen placed France among the 'conservative' regimes based on multiple welfare program features including public pension structures, we argue as do others (Hantrais 2004; Pailhé, Rossier, and Toulemon 2008) that its large-scale, publicly-subsidized day-care integrated with maternity leave puts France much more among the 'dual-earner' regimes specifically in the family-policy domain of welfare regimes. We use the term 'dual-earner' and not Esping-Andersen's "social democratic" term both to allow France to be included in the same group as Norway, and to emphasize the gender dimension of regime type. In ranking France similarly to the Nordic countries, Gornick et al (1997) and Pampel (2001) accordingly characterize the regime character with "employment support for mothers" and "women-friendliness" scales.

The use of regime-type classifications of family policies is a key feature of our study. Estimating the effects of single family-policy measures does not address the argument that the effects of family policies are best understood as an interacting and, in the best case, mutually-supporting family-policy "regime" (Brewster and Rindfuss 2000; Thévenon and Gauthier 2011). The data demands for obtaining comparable measures across countries on potentially interacting policy measures such as maternity leave provisions and subsidized childcare are challenging, notwithstanding improvements due to the recent assembly of cross-national

databases on these (e.g., Thévenon 2012). Empirically, where both family-policy regime type and specific measures have been included as predictors, regime type has been found to be a much stronger predictor of fertility than have specific measures (Del Boca et al 2009). Salles, Rossier, and Brachet (2010) argue additionally that policy support for the combining of employment and motherhood can generate cultural shifts in attitudes towards mothers' remaining in the labor force, and this is again an effect that is more of the nature of a regime effect than of variation in a specific family-policy provision at any one time and place. Such cultural shifts may also narrow the differences between legal provisions (e.g., for maternity leave) and their implementation by employers.

Panel data estimation

Unlike previous studies of socio-economic differentials in fertility and family-demographic events, in which retrospective fertility and marital and cohabitation histories have been typically relied on (e.g., Billari and Philapov 2004; Martin 2000, 2006; Perelli-Harris et al 2010; 2012), our study relies exclusively on panel data. The countries of our study are accordingly chosen from among others with similar family-policy regimes due to their having panel surveys of sufficiently long, annual periodicity and with distinctions in their data between marital and non-marital cohabiting unions. We use the Panel Study of Income Dynamics (PSID) for the United States, the Household Income and Labor Dynamics in Australia (HILDA), the German Socio-economic Panel (SOEP), and the Swiss Household Panel (SHP), and the French and Norwegian European Union Survey of Income and Living Conditions (EU SILC).

The Cornell Cross-National Equivalent File (CNEF, Frick et al 2007) has coded the PSID, HILDA, SOEP, and SHP data into harmonized formats. The CNEF versions come with the significant advantage of having already been pre-coded to have employment, schooling, partnership, and household and family structure variables with consistent variable names and variable values across the countries. We use these CNEF versions of the HILDA, SOEP, and

SHP data, and code exposure in single-year intervals from 2001-02, 2002-03, through 2009-10. For the PSID, which has a panel interval of two years, we use the in-country (U.S.) data file in combination with the CNEF so that we are able to code the birth event in a one-year interval. This is the calendar year between the panel years. This means that we have exposure in the PSID only every other year, 1998, 2000, though 2008.

The CNEF group of panel surveys includes no country from the "two-earner" family-policy regime type. To include countries from this group with comparable panel survey data, we use the seven to eight available panel years of the French and Norwegian European Union Survey of Income and Living Conditions (EU SILC), from 2004 to 2010 in France and from 2004 to 2011 in Norway. These are the only two EU SILC countries for which the panel period is more than four years (Eurostat 2011). We build on our previous use of the in-country versions of the French (Toulemon and Pennec 2010) and Norwegian (Kitterød, Rønsen, and Seierstad forthcoming) EU SILC data sources.

Our panel data have three advantages. First and foremost, panel data allow us to use both men's and women's employment data as predictors of first parenthood. Second, unlike the major cross-nationally harmonized demographic surveys used to analyze differences in fertility, the Fertility and Family Surveys (FFS) of the 1990s (see, for example, Billari and Philapov 2004), the European Community Household Panel (ECHP) of the 1990s (see especially Adsera 2004, 2005), and the Gender and Generations Survey (GGS) of the 2000s (see, for example, Perelli-Harris et al 2012), our collection of panel surveys allows us not only to include two countries from the 'liberal' group, Australia and the U.S. Previous analyses of fertility rates have shown that the English-speaking countries have common age-specific patterns (Chandola, Coleman, and Hiorns 2002) that are not found in developed countries outside of the 'liberal' group (Chandola et al 1999), and a common pattern of early first births among less educated women (Sullivan 2005; Sigle-Rushton 2008; Rendall et al 2010). Including Australia and the U.S. allows us to distinguish two "family-unfriendly" regime types, 'liberal' and 'conservative.'

Third, reporting biases in retrospective data are much reduced in panel data. In a primarily low-income sample, Teitler, Reichman, and Koball (2006) found women's retrospective reporting of births within or outside of cohabitation to be systematically biased towards successful cohabitation relationships, and found these biases so severe as to call into question the validity of using retrospective reports of cohabitation. Peters (1988) found retrospective reports of dates of marriage to be less consistent with panel-observed marriage for women with lower educational attainment. Rendall et al (1999) found men's fertility estimates to be strongly downwardly biased in retrospective reports but not in panel reports (see also Joyner et al 2012).

Panel data, however, have three significant disadvantages relative to retrospective data. First, fewer years of exposure are obtained from panel than retrospective data. Second, attrition in panel studies may be substantial and may bias estimates in ways that do not occur in retrospective data. Third, birth, partnership, and employment histories are left-censored in the year in which the panel begins. The first and third disadvantages imply a trade-off between the advantages of including as many years of exposure to first birth and of including an employment trajectory among the predictor variables. Because the HILDA, SHP, and French and Norwegian EU SILC panels all began since 2001, the longer the panel employment trajectory that is used as a predictor of entry to parenthood, the fewer are the available years for the first-parenthood outcome variable. For this reason, we use only employment in the year before exposure to first parenthood.

The first-parenthood hazard

We model first parenthood as a binary outcome over a 12-month period. This 12-month period is between survey waves one year apart for the EU SILC, HILDA, SHP, and SOEP, and for the calendar year immediately after the previous survey wave in the PSID. For all except the U.S. and France, we rely on a variable for the presence of any minor-aged children in the household to identify parity-zero women, and to identify men partnered to a parity-zero women. For this

reason, we limit our analyses to 18 to 34 year old women and to men partnered to women in this age range. For women 35 and over, a first-born child may be old enough to have already left the household.

Formally, we specify entry to first-parenthood as a discrete-hazard process characterized by exposure to the transition from not being a parent at age *a*-1 in the previous year to being a parent for the first time at age *a*. We represent this fertility hazard *f* for having a first birth b_1 between ages *a-1* and *a* by $\boldsymbol{f(b_1; a)}$. The predictor variables that characterize the type of "exposure" to the hazard of first birth (parenthood) represent the characteristics of individuals of reproductive ages.

We therefore estimate three **first-parenthood hazards**: two first-parenthood hazard equations for women, $f(b_{1,u}; a)$ for **unpartnered** and $f(b_{1,p}; a)$ for **partnered women**, and one first-parenthood hazard equation $f(b_{1,p}; a)$ for **partnered men**. For men, we ignore unpartnered (non-coresident) fertility due to data limitations across our six countries, and in the present version of the paper we have not included results for partnered men in France. The hazard's predictor variables include own employment status *e* and educational attainment *s*. Partnership status, educational attainment, and employment status *e* are all defined at age *a-1*, the year immediately before this age-*a* exposure to first parenthood. The functional form we use throughout is:

$$Prob(b; a, s, e) = Logit[\,\beta_0 + \beta_1(a-25) + \beta_2(a-25)^2 + \boldsymbol{\beta_3 s} + \boldsymbol{\beta_4 e}]$$

Employment status *e* is operationalized through the aggregate number of employed hours in the last year, divided into categories of < 750, 750-1,749, and 1,750+ hours. We use 1,750+ hours as the reference category, which we describe "fulltime, full-year employment". We categorize 750-1,749 as "part-time or part-year employment" and < 750 as "little or no employment". A number of studies have examined the relationship of employment to first births, and to fertility

more generally, in the countries under study, and have used more detailed characterizations of employment that include more than one year of employment or that differentiate between fixed-term and permanent employment, especially with respect to Germany (Ozcan, Mayer, and Luedicke 2010; Kreyenfeld 2009; Schmitt 2012; Tolke and Diewald 2003). We are unable to do similarly here in cross-nationally comparative ways, or without losing substantial numbers of years of fertility exposure in most of our countries.

We would have liked to estimate models that interact employment with educational attainment, allowing us to test predictions of more homogeneous effects of employment by SES in 'dual-earner' than 'liberal' or 'conservative' regimes. We were limited by sample sizes in our ability to estimate such interactions across each of the six countries. We therefore included educational attainment only as a control variable. We coded educational attainment, in the context of large difference in educational systems across countries, based on obtaining sufficient numbers of individuals in each educational attainment category in each country. We coded two categories of secondary school educational attainment in France and Norway, depending on whether the highest secondary qualification was obtained. We code two categories of tertiary education in the U.S., depending on whether a bachelors degree (4 years of university education) was obtained. For Australia, Germany, and Switzerland, we coded only whether any tertiary qualification was obtained.

The different countries' sample designs lead to differences in the oversampling of groups whose fertility patterns are likely to differ from other groups. Our regression models, moreover, use relatively few predictor variables. We therefore apply sample weights in both our descriptive and multivariate analyses.

Study Hypotheses

Family-policy regimes differ greatly in their support for combining employment and childrearing. Accordingly, we expect that different family-policy regimes and social norms will create different

gender inequalities in the relationship of employment to first birth. We expect these gender inequalities to be substantial in both ‘conservative’ and ‘liberal’ regimes, and relatively weak in ‘dual-earner’ regimes.

The following hypotheses are tested:

H1: More employment *reduces* **partnered women’s** entry to parenthood in ‘liberal’ and ‘conservative’ regimes and *increases* it in ‘dual-earner’ regimes.

H2: More employment *increases* **partnered men’s** hazards of entry to parenthood in all regimes (‘liberal’, ‘conservative’, and ‘dual-earner’).

We are also interested in capturing first births out of unions. The means-testing institutional feature of ‘liberal’ regimes is expected to facilitate childrearing among women with little or no employment. We therefore propose the following hypothesis that is complementary to hypothesis H1:

H3: More employment *reduces* **unpartnered women’s** first birth hazards in ‘liberal’ regimes.

Only in ‘liberal’ regimes is unpartnered childbearing a common phenomenon. We expect unpartnered first-birth hazards to be too low in ‘dual-earner’ and ‘conservative’ regimes for there to be any discernible effects of employment on first parenthood. Additionally, we are largely unable to identify first-time fathers outside a union in our panel surveys, and therefore we construct no hypotheses about regime effects of employment of unpartnered men on their first fatherhood.

RESULTS

In our results, we separate partnered and unpartnered women age 18-34, and analyze partnered women separately from men partnered to 18-34 year old women. We first describe

our samples of unpartnered and partnered women and partnered men (see Table 1). Around two-thirds of women's 18-34 year-old years exposed to first birth are unpartnered across our six countries, but with large variations. In five of our six countries, the range is from a low of 60% unpartnered in Australia to a high of 71% unpartnered in the United States. However, Germany is highly exceptional with 88% unpartnered. In part this may be due the very high rate of entry to motherhood in Germany after becoming married or entering a cohabiting union, as seen in Table 1 by partnered German women having the highest annual proportion entering motherhood (21.6%).

Notable also are the older median ages of partnered women and men with no children in the 'conservative' countries, at 30 and 33 in Germany and 29 and 33 in Switzerland. The median ages are substantially lower in the 'liberal' regimes of Australia (27 and 30) and the U.S. (28 and 30). The median ages of partnered women and men with no children are lower still, however, in the "dual earner" countries, respectively 25 and 28 for partnered women and men in Norway and 27 for partnered women in France. The older median ages of the conservative countries will have been produced by some combination of historical fertility declines that have reduced successive cohort sizes and later entry to unions and later entry to parenthood after entering a union. Although we control for age in our analyses of the multivariate associations of employed-hours with first parenthood, our interpretations nevertheless need to take into account the endogeneity of age in the process we are modeling. In particular, Germany's partnered women may be a highly selected group.

[TABLE 1 ABOUT HERE]

Comparing next educational attainment and employed-hours distributions across countries, it is notable that comparisons are easier to make with respect to employed-hours than with respect to educational attainment. Differences in educational systems make it difficult to

differentiate between two secondary school levels (lower and upper secondary) in Australia and the U.S., but we are able to differentiate between two tertiary school levels in the U.S. and between two secondary school levels in France and Norway. The six countries exhibit a relatively large range around approximately 50% tertiary educated. The U.S. and Germany were the only two countries in which more than 50% of unpartnered women were tertiary educated. The lowest levels of education were those for Swiss women, with only 18.5% of unpartnered women and 32.7% of partnered women with a tertiary education.

We are able to distinguish three categories of hours worked in the previous year consistently across all six countries. Of particular interest for our analyses of gender inequalities are differences between the employed-hours distributions of partnered women and partnered men. Remarkable here are the consistently much greater proportions of partnered men than of partnered women employed for a full-time, full-year level of aggregate hours (1,750+, or the equivalent of at least a 35 hour week for 50 weeks). Perhaps somewhat surprisingly given the absence of any children in the household, partnered men's full-time, full-year percentages exceeded partnered women's by substantial amounts in all six countries. In France and Norway, just under half of partnered women worked full-time, full-year, and only 56.8% did in Germany. Partnered men's full-time, full-year percentages exceeded partnered women's by the largest amount, 26.6 percentage points, in Germany, and the least, 14.8 percentage points, in Switzerland. Only in the U.S., however, do unpartnered women's percentages in full-time, full-year employment in the preceding year match those of partnered women (61.4% of unpartnered versus 60.4% of partnered women without children).

[TABLES 2A, 2B, AND 2C ABOUT HERE]

We present in Tables 2a, 2b, and 2c our regression model estimates. We find partial support for our first hypothesis, that more employment *reduces* **partnered women's** 1st birth

hazards in ‘liberal’ and ‘conservative’ regimes and *increases* it in ‘dual-earner’ regimes (Table 2b). Consistent with this hypothesis, in the ‘dual-earner’ regimes of France and Norway, < 750 hours of employment is associated with *lower* 1st birth hazards compared to full-time, full-year employment. The magnitude of this coefficient is higher in Norway (-0.660) than in France (-0.295). In France, women in part-time or part-year employment have the highest likelihood of entry to first motherhood. Part-time employment is less common in France, however, than in Norway (see again Table 1). Also noteworthy is that in both France and Norway, the negative employment coefficient is offset by a negative education coefficient (relative to the lower-secondary category) of almost equal strength, although only in France is this statistically significant. The choice of reference category is important for education, however, since only 15% and 16% respectively of French and Norwegian partnered women fall into this lowest education reference category (see again Table 1).

In neither of the ‘liberal’ regimes of Australia and the U.S. is the < 750 hours coefficient for partnered women’s first birth hazard significantly different from zero. In Australia, the positive coefficient of 0.407 for ‘part-time or part-year employment’ status relative to full-time, full-year employed is consistent with the prediction that more employment is associated with a lower first-birth hazard in ‘liberal’ and ‘conservative’ regimes.

The employment coefficients for partnered women in Germany, a ‘conservative’ regime, are unexpectedly in the negative direction relative to the reference category full-time, full-year employment. Moreover, with a magnitude of -0.787 the coefficient for ‘little-or-no work’ (< 750 hours) is similar to that for partnered women in Norway (-0.744). Unfortunately the CNEF does not include variables that allow us to distinguish reasons for working few or no hours, and therefore we are not able to compare housewives’ to students’ and unemployed women’s first birth propensities between these countries.

We find relatively weak support for our second hypothesis that employment *increases* **partnered men’s** first birth hazards in ‘liberal’, ‘conservative’, and ‘dual-earner’ regimes (Table

2c). Only in Norway is the coefficient for ‘little or no work’ (-0.608) negative and statistically significant at the $p < .05$ level relative to the reference full-time, full-year employment category. Only in the U.S. is there an indication, moreover, of a positive income effect being picked up instead through the educational attainment variables, with the coefficient of 0.752 for at least a bachelors degree being positive and significant at $p < .01$. Evidence for relatively little dependence on men’s employment and income in Australia is seen both in the marginally negative coefficient for men being tertiary-educated coefficient (-0.201, p=0.062) relative to secondary-educated only and in the marginally positive coefficient (0.279, p=0.094) for being part-time/part-year employed relative to full-time/full-year employed --- that is, *less* male employment is associated with higher rates of entry to first parenthood.

We present **unpartnered women’s** first birth equations for only four of our six countries, omitting Norway and Switzerland, for which the annual first birth hazard was below 1% (see again Table 1) and therefore too small to obtain sufficient sample sizes of unpartnered women with a first birth to be able to estimate the socio-economic associations. In the ‘liberal’ countries, we did find a statistically-significant positive coefficient for ‘little-to-no’ (< 750 hours) employment in Australia (0.575, p<.05), consistent with our third hypothesis, but found no statistically-significant findings for the U.S. For Germany, for which we had not made a prediction as to the direction of effect, a marginally-significant negative coefficient is found (-0.280, p=0.090), suggesting that both partnered *and* unpartnered women with ‘little-to-no employment’ in Germany are less likely to enter motherhood than are partnered and unpartnered women with full-time/full-year employment. In France, unpartnered women with ‘little or no employment’ are less likely to enter motherhood than are unpartnered women with full-time/full-year employment.

DISCUSSION

We presented new evidence on entry to first parenthood at ages 18 to 34 across six countries representing three family-policy “regimes.” The unique empirical contribution of our study was to

estimate identical models of current individual employment on women's and men's entry to first parenthood using comparable panel data covering approximately the 2000s decade. Our overarching research question was whether the relationship between employment and family formation differs systematically across family-policy and macro-social contexts. We were especially interested in testing hypotheses generated from theory of "reproductive polarization", in which family policy is claimed to play a central role in generating or mitigating socio-economic heterogeneity in family formation. A well-integrated set of family-policy provisions that together form a 'dual-earner' family-policy regime is expected to provide strong incentives for women across socio-economic levels to first secure permanent employment before entering parenthood, and then to proceed quickly to childbearing with their jobs held for their return after each maternity leave (Gustafsson 2001). This is expected to generate a strong dependence of first birth on being employed among women in those regimes. Countries that do not provide such a set of policies have been grouped into two categories, 'liberal' and 'conservative' regimes (Esping-Andersen 1999). Overall fertility rates have been found to be much higher in 'liberal' than 'conservative' regimes, and the patterns of fertility in 'conservative' regimes have been broadly characterized as being marked by late childbearing and high rates of childlessness (Sánchez-Barricarte and Fernández-Carro 2007), whereas in 'liberal' regimes the phenomenon of early childbearing among socio-economically disadvantaged women has been considered a defining characteristic (McLanahan 2004; Chandola et al 2002). Considerable heterogeneity of strategies for combining career paths of employment and childrearing are also expected in 'liberal' regimes (Joshi 2002). We tested whether and how these two alternative regime types to the 'dual-earner' type generate gender inequalities in the relationship of employment to first parenthood.

Our findings were mostly either in support of our hypotheses or were statistically non-significant, with Germany, however, providing some findings opposite to those hypothesized. A first note is our finding of surprisingly large and pervasive gender inequality on our main

predictor variable, employed hours in the previous year among partnered individuals with no children. We saw a range of between 15 and 25 percentage-point excesses of men's over women's full-year, full-time employed proportions in the year *before* exposure to entry to first parenthood. No particular pattern between regime types, moreover, was apparent, with two 'conservative' countries, Switzerland and Germany, providing both the lower and upper bounds respectively of this range. These employment differences were not obviously related to gender differences in educational attainment, a finding that reinforces the value of our having extended previous comparative family-policy-regime analyses of first births by education (e.g., Rendall et al 2010) to consider employment as a predictor variable. As expected, however, smaller numbers of observations of first-parenthood events in our panel data (compared to retrospective data on fertility) led to many statistically non-significant regression coefficient estimates, and prevented us from being able to model theoretically-interesting interactions of employment and education on first parenthood. Neither, moreover, did we attempt to estimate models that simultaneously considered the effects male and female partners' employment on first partnered parenthood, nor that distinguished between durations of partnership or between the married versus cohabiting status of the couple.

In the 'dual-earner' regime group, Norwegian and French findings from the regression analyses were largely in line with predictions. 'Little or no work' in the previous year was associated with a lower likelihood of first birth among both partnered women and (in Norway) partnered men, providing support for both our first and second hypotheses. Further support for our first hypothesis was seen from the 'liberal' regime group, in which Australian women with part-time/part-year work in the previous year had a higher likelihood of first birth among partnered women than did otherwise similar partnered Australian women in full-time, full-year work. Against our study's first hypotheses, however, was the finding for partnered German women that full-time, full-year work in the previous year was associated with a higher likelihood of first birth.

Regarding again our study's second hypothesis, this was supported in both Germany and Norway, where 'full-time, full-year' work in the previous year was associated with a statistically higher likelihood of first birth than 'little or no work' among partnered men. In Switzerland and the U.S., however, no statistically-significant partnered male employment effect was detected. Among partnered men in Australia, 'part-time or part-year work' in the previous year was associated with a higher likelihood of first entry to parenthood than was full-time, full-year work, although this coefficient was only marginally significant.

Finally, concerning our third hypothesis, 'little or no work' in the previous year was associated with the highest likelihood of first birth among unpartnered women in the 'liberal' regime of Australia, thus providing some support for the hypothesis. Also supportive of the particularity of the inverse relationship between employment and entry to motherhood in 'liberal' regimes were the findings among unpartnered women in France and in Germany of a lower likelihood of first birth after a year of 'little or no work' than after a year of 'full-time, full-year' work. We had not expected to find sufficient cases to discern an effect of employment on unpartnered first births in other than the 'liberal' regimes, and it should be kept in mind that our definition of unpartnered is as at the year *before* exposure to first parenthood. Many of these first births in France and Germany may have occurred after becoming partnered. Even in Australia, we found that about for about half of first births to women who were unpartnered before the year of fertility exposure, the woman was partnered at the end of the year of fertility exposure (results not shown).

Evidence of very high unpartnered proportions of 18-34 year old parity-zero women in Germany was also striking, indicating a possible alternate response in an environment unfavorable for combining employment and motherhood: that of delaying or foregoing entry to marriage or cohabitation as a means of delaying or foregoing entry to motherhood. Germany is a country for which there are also several alternate estimates of employment effects on entry to parenthood (Gonzalez and Jurado-Guerrero 2006; Ozcan, Mayer, and Luedicke 2010;

Kreyenfeld 2009; Schmitt 2012; Tolke and Diewald 2003), in most cases from the original data source, the German Socioeconomic Panel, from which our harmonized CNEF data were derived. A future task before us for both Germany and our other countries is to reconcile our results with previous findings. Schmitt (2012), for example, in examining women's first births in the period 1991-2007, noted both a relatively fast entry to first parenthood after first employment, consistent with our results, and also substantial proportions of women entering parenthood before having entered employment, contrary to our overall finding of a positive effect of employment on first motherhood.

The large overall gender inequalities in employed hours across our six countries, the German results of very high proportions of unpartnered women aged 18-34 without children, and the older ages of partnered women and men in the 'conservative' countries than in the 'dual-earner' countries, all point to the need to understand better the processes of selection into the conditions from which our regression analyses begin. In an earlier study using 1994-2001 panel data from the European Community Household Panel (ECHP), Gonzalez and Jurado-Guerrero (2006) modeled selection into childlessness at the first wave as a prior process to that of first birth by employment and partnership status. Selection into partnership status and employment may be additional processes that merit formal modeling. Moreover, these processes of employment and partnership formation (and dissolution) are known to vary by indicators of SES including educational attainment. Therefore examination of these selection processes in combination with the process of entry to first parenthood conditional on employment and partnership status may provide greater insights into the operation of reproductive polarization across 'liberal' and 'conservative' regimes, and into the relative absence of such polarization in 'dual-earner' regimes.

REFERENCES

Aassve, A., F.C. Billari, and Z. Spéder (2006) “Societal transition, policy changes and family formation: Evidence from Hungary.” *European Journal of Population* 22:127-152.

Adsera, A. (2004) **“**Changing fertility rates in developed countries: The impact of labor market institutions.” *Journal of Population Economics* 17:17-43.

Adsera, A. (2005) “Vanishing Children: From High Unemployment to Low Fertility in Developed Countries.” *American Economic Review Proceedings* 95(2):189-193.

Billari, F.C., and D. Philapov, (2004) “Education and the Transition to Motherhood: A Comparative Analysis of Western Europe” *European Demographic Research Papers 3.*

Brewster, K. L., and R. R. Rindfuss (2000) Fertility and women’s employment in industrialized nations *Annual Review of Sociology* 26:271-296.

Callens, M., and C. Croux (2005) “The impact of education on third births: A multilevel discrete-time hazard analysis.” *Journal of Applied Statistics* 32(10):1035-1050.

Castles, F.G. (2003) “The world turned upside down: Below-replacement fertility, changing preferences and family-friendly public policy in 21 OECD countries.” *Journal of European Social Policy* 13(3):209-227.

Chandola, T., D. A. Coleman, and R. W. Hiorns (1999) “Recent European fertility patterns: Fitting curves to ‘distorted’ distributions.” *Population Studies* 53:317-329.

Chandola, T., D. A. Coleman, and R. W. Hiorns (2002) “Distinctive features of age-specific fertility profiles in the English-speaking world: Common patterns in Australia, Canada, New Zealand and the United States, 1970-98.” *Population Studies* 53:317-329.

Cherlin, A.J. (2010) “Demographic Trends in the United States: A Review of Research in the 2000s.” *Journal of Marriage and Family* 72:403-419.

Del Boca, D., S. Pasqua, and C. Pronzato (2009) “Motherhood and market work decisions in institutional context: a European perspective.” *Oxford Economic Papers* 61:147–171.

DiPrete, T.A., S.P. Morgan, H. Engelhardt, and H. Pacalova (2003) “Do cross-national differences in the costs of children generate cross-national differences in fertility rates?” *Population Research and Policy Review* 22:439-477.

Dorbritz, J. (2008) “Germany: Family diversity with low actual and desired fertility” *Demographic Research* 19:557-598

Duncan, G.J., and J. Brooks-Gunn (Eds.) (1997) *Consequences of Growing Up Poor* New York: Russell Sage Foundation.

Ekert-Jaffé, O., H. Joshi, K. Lynch, R. Mougin, and M. S. Rendall (2002) “Fertility, Timing of Births, and Socio-economic Status in France and Britain: Social Policies and Occupational Polarisation.” *Population-E* no.3:475-508.

Ermisch, J. F. (1989) “Purchased child care, optimal family size and mother’s employment.” *Journal of Population Economics* 2(2):79-102.

Esping-Andersen, G. (1990) *The Three Worlds of Welfare Capitalism* Princeton: Princeton University Press.

Esping-Andersen, G. (1999) *Social Foundations of Post-industrial Economies* Oxford: Oxford University Press.

Eurostat (2011) “2008 Comparative EU Final Quality Report,” Version 3, July 2011.

Frejka, T., and J.P. Sardon (2007) “Cohort birth order, parity progression ratio and parity distribution trends in developed countries.” *Demographic Research* 16:315-374.

Frick, J.R., S.P. Jenkins, D.R. Lillard, O. Lipps, M. Wooden (2007) “The Cross-National Equivalent File (CNEF) and its Member Country Household Panel Studies.” *Schmollers Jahrbuch* 127:627-654.

Gauthier, A. H. (2007) “The impact of family policies on fertility in industrialized countries: A review of the literature” *Population Research and Policy Review* 26(3):323-346.

Goldstein, J.R., and C.T. Kenney (2001) “Marriage Delayed or Marriage Forgone? New Cohort Forecasts of First Marriage for U.S. Women.” *American Sociological Review* 66(4):506-19.

Gonzalez, M.-J. and Jurado-Guerrero, T. 2006. “Remaining childless in affluent economies: A comparison of France, West Germany, Italy and Spain, 1994-2001. “ *European Journal of Population* 22: 317-352.

Gornick, J.C., M.K. Meyers, and K.E. Ross (1997) “Supporting the employment of mothers: Policy variation across fourteen welfare states.” *Journal of European Social Policy* 7(1):45-70.

Gottschalk, P., and S. Danziger (1993) “Family structure, family size, and family income: Accounting for changes in the economic well-being of children, 1968-1986.” In S. Danziger and P. Gottschalk (eds) *Uneven Tides: Rising Inequality in America* New York: Russell Sage

Gustafsson, S.S. (2001) “Optimal age at motherhood: Theoretical and empirical considerations in postponement of maternity” *Journal of Population Economics* 14:225-247.

Gustafsson, S.S., and C. Wetzels (2000) “Optimal age for first birth: Germany, Great Britain, the Netherlands and Sweden” in Gustafsson, S. S., and D. E. Meulders (Eds) *Gender and the Labour Market: Econometric Evidence of Obstacles to Achieving Gender Equality* New York: St. Martin’s.

Gustafsson, S.S., C.M.M.P. Wetzels, J.D. Vlasblom, and S. Dex (1996) “Women’s labor force transitions in connection with childbirth: A panel data comparison between Germany, Sweden and Great Britain.” *Journal of Population Economics* 9(3):223-246.

Hantrais, L. (2004) *Family Policy Matters: Responding to Family Change in Europe* Bristol: The Policy Press.

Harknett, K, and A. Kuperberg (2011) “Education, Labor Markets and the Retreat from Marriage” *Social Forces* 90(1):41-64.

Heard, G. (2011) “Socioeconomic Marriage Differentials in Australia and New Zealand.” *Population and Development Review* 37(1):125-160.

Henderson, A., and L.A. White (2004) “Shrinking welfare states? Comparing maternity leave benefits and child care programs in European Union and North American welfare states 1985-2000” *Journal of European Public Policy* 11(3):497-519.

Heuveline, P., and Timberlake (2004) “The Role of Cohabitation in Family Formation: The United States in Comparative Perspective” *Journal of Marriage and Family* 66:1214-1230.

Hoem, J.M., and M. Kreyenfeld (2006) “Anticipatory analysis and its alternatives in life-course research. Part 1: The role of education in the study of first childbearing.” *Demographic Research* 15(16):461-484.

Hoffman, S.D., and E.M. Foster (2000) “AFDC benefits and nonmarital births to young women.” *Journal of Human Resources* 35(2):376-391.

Joshi, H. (2002) Production, Reproduction, and Education: Women, Children, and Work in a British Perspective *Population and Development Review* 28(3): 445-474.

Joyner K., H.E. Peters, K. Hynes, A. Sikora, J.R. Taber, and M.S. Rendall (2012) “The quality of male fertility data in major U.S. surveys.” *Demography* 49(1):101-124.

Kitterød, R.H., M. Rønsen, and A. Seierstad (forthcoming) “Mobilizing female labour market reserves: What promotes women’s transitions between part-time and full-time work?” *Acta Sociologica* 56(2).

Kohler, H.P., F.C. Billari, and J.A. Ortega (2002) “The emergence of lowest-low fertility in Europe during the 1990s” *Population and Development Review* 28(4):641-680.

Kohler, H.P., and J.A. Ortega (2002) “Tempo-Adjusted Period Parity Progression Measures: Assessing the Implications of Delayed Childbearing for Cohort Fertility in Sweden, the Netherlands and Spain” *Demographic Research* 6:145-190.

Kravdal Ø. (2004) “The Impact of Individual and Aggregate Unemployment on Fertility in Norway” *Demographic Research* 6:263-294.

Kreyenfeld, M. (2009). “Uncertainties in Female Employment Careers and the Postponement of Parenthood in Germany.” *European Sociological Review* 26(3): 351-366.

Lappegård, T. and M. Rønsen (2005) “The Multifaceted Impact of Education on Entry into Motherhood.” *European Journal of Population* 21:31-49.

Lappegård, T and M Rønsen (forthcoming) “Socioeconomic differentials in multipartner fertility.” *Demography*

Lesthaeghe, R. (2010) The unfolding story of the second demographic transition *Population and Development Review* 36:(2)211-251.

Manning, W.D., and S. Brown (2006) “Children’s economic well-being in married and cohabitating parent families.” *Journal of Marriage and the Family* 68:345-362.

Martin, S.P. (2000): “Diverging fertility among U.S. women who delay childbearing past age 30.” *Demography* 37: 523-533.

Martin, S.P. (2006) “Trends in Marital Dissolution by Women’s Education in the United States.” *Demographic Research* 15:537-560.

Matysiak, A., and D. Vignoli (2008) “Fertility and Women’s Employment: A Meta-analysis.” *European Journal of Population* 24:363-384.

McDonald, P. (2006) “Low fertility and the state: The efficacy of policy” *Population and Development Review* 32(3):485-439.

McLanahan, S. (2004) “Diverging destinies: How children fare under the second demographic transition.” *Demography* 41(4):607-627.

Meron, M., and I. Widmer (2002) “Les femmes au chomage retardent l’arrivée du premier enfant [Unemployment leads women to postpone the birth of their first child]” *Population* 57(2):327-358.

Morgan, S.P. (2003) Low fertility in the twenty-first century *Demography* 40(4):589-603.

Morgan, S. P., E. Cumberworth, and C. Wimer (2011) “The Great Recession’s Influence on Fertility, Marriage, Divorce, and Cohabitation.” Pp.220-246 in D. Grusky, B. Western, and C. Wimer (Eds.) *The Great Recession* New York: Russell Sage.

Ni Bhrolchain, M., and E. Beaujouan (2012) “Fertility postponement is largely due to rising educational enrolment.” *Population Studies*

Ozcan, B., Mayer, K. U., and J. Luedicke (2010) “The Impact of Unemployment on the Transition to Parenthood.” *Demographic Research* 29: 807-846.

Pailhé, A., C. Rossier, and L. Toulemon (2008) “French family policy: Long tradition and diversified measures.” *Vienna Yearbook of Population Research* 149-164.

Pailhé, A., and A. Solaz (2012) “The influence of employment uncertainty on childbearing in France: A tempo or quantum effect?” *Demographic Research* 26-1:1-40.

Pampel F. (2001) *The Institutional Context of Population Change: Patterns of Fertility and Mortality across High-Income Nations* Chicago: University of Chicago Press.

Panel Study of Income Dynamics (2009) “Public Release Family File Codebook.” Produced and distributed by the Institute for Social Research, Survey Research Center, University of Michigan, Ann Arbor, MI 2013.

Panel Study of Income Dynamics (no date) “Interviewing Procedures.” Produced and distributed by the Institute for Social Research, Survey Research Center, University of Michigan, Ann Arbor, MI 2013.

Perelli-Harris, B., et al (2010) “The Educational Gradient of Childbearing Within Cohabitation in Europe.” *Population and Development Review* 36(4):775-801.

Perelli-Harris, B., M. Kreyenfeld, W. Sigle-Rushton, R. Keizer, T. Lappegard, A. Jasilioniene, C. Berghammer, and P. Di Giulio (2012) “Changes in union status during the transition to parenthood: An examination of 11 European countries.” *Population Studies* 66(2): 167-182.

Peters, H.E. (1988) "Retrospective Versus Panel Data in Analyzing Life Cycle Events." *Journal of Human Resources* 23:488-513.

Rendall, M.S., L. Clarke, H.E. Peters, N. Ranjit, and G. Verropoulou (1999) “Incomplete reporting of male fertility in the United States and Britain: A research note.” *Demography* 36(1):135-144.

Rendall, M.S., E. Aracil, C. Bagavos, C. Couet, A. DeRose, P. DiGiulio, T. Lappegard, I. Robert-Bobée, M. Rønsen, S. Smallwood, and G. Verropoulou (2010) “Increasingly heterogeneous ages at first birth by education in Southern-European and Anglo-American family-policy regimes: A seven-country comparison” *Population Studies* 64(3):209-227.

Rendall, M.S., O. Ekert-Jaffé, H. Joshi, K. Lynch, and R. Mougin (2009) “Universal versus economically polarized change in age at first birth: a French-British comparison” *Population and Development Review* 35(1):89-116.

Rønsen M and K Skrede (2008) “Fertility trends and differentials in the Nordic countries: Footprint of welfare policies and challenges on the road ahead.” *Vienna Yearbook of Population Research* 2008: 103-123

Rønsen, M. and M. Sundström (2002) “Family policy and after-birth employment among new mothers – A comparison of Finland, Norway and Sweden.” *European Journal of Population* 18:121-152.

Rosenzweig, M.R. (1999) “Welfare, marital prospects, and nonmarital childbearing.” *Journal of Political Economy* 107(6):S2-S32.

Ruhm, C. (1998) “The economic consequences of parental leave mandates: Lessons from Europe.” *Quarterly Journal of Economics* 113(1):285-317.

Salles, A., C. Rossier, and S. Brachet (2010) Understanding the long term effects of family policies on fertility: The diffusion of different family models in France and Germany *Demographic Research* 22(34):1057-1096.

Sánchez-Barricarte, J.J., and R. Fernández-Carro (2007) Patterns in the delay and recovery of fertility in Europe *European Journal of Population* 23:145-170.

Schmitt, C. (2012) “Labour market integration, occupational uncertainties, and fertility choices in Germany and the UK” *Demographic Research* 26:253-92

Schulze, H., and H. Tyrell (2002) “What happened to the European family in the 1980s? The polarization between the family and other forms of private life.” Ch.3 (pp.69-119) in F. Kaufmann, A. Kuijsten, H. Schulze, and K.P. Strohmeier (eds) *Family Life and Family Policies in Europe, Volume 2, Problems and Issues in Comparative Perspective* Oxford: Oxford University Press.

Shang, Q., and B.A. Weinberg (2013) “Opting for families: Recent trends in the fertility of highly educated women” *Journal of Population Economics* 26:5-32.

Sigle-Rushton, W. (2008) “England and Wales: Stable Fertility and Pronounced Social Status Differences.” *Demographic Research* 19: 459-502.

Smeeding, T., I. Garfinkel, and R.B. Mincy (2011) “Young Disadvantaged Men: Fathers, Families, Poverty, and Policy.” *The Annals of the American Academy of Political and Social Science* 635:6-21.

Sobotka, T., and V. Skirbekk, and D. Philipov (2011) “Economic Recession and Fertility in the Developed World” *Population and Development Review* 37(2):267-306.

Stier, H., N. Lewin-Epstein, and M. Braun (2001) “Welfare regimes, family-supportive policies, and women’s employment along the life-course.” *American Journal of Sociology* 106(6):1731-60.

Sullivan, R. (2005) “The age pattern of first-birth rates among U.S. women: The bimodal 1990s.” *Demography* 42(2):259-273.

Teitler, J., N.E. Reichman, and H. Koball (2006) “Contemporaneous Versus Retrospective Reports of Cohabitation in the Fragile Families Survey.” *Journal of Marriage and Family* 68(2):469–477.

Thévenon, O. (2011) “Family Policies in OECD Countries: A Comparative Analysis.” *Population and Development Review* 37(1):57-87.

Thévenon, O. (2012) “Macro-level database on fertility and policies supporting families with children in European and OECD countries.” Reproductive Decision Making in Macro-Micro Perspective (REPRO) project report, Paris: INED.

Thévenon, O., and A. Gauthier (2011) “Family Policies in Developed Countries: A ‘Fertility booster’ with Side-Effects.” *Community, Work and Family* 14(2):197-216.

Thévenon, O. and A. Solaz (2013) “Labour market effects of parental leave policies in OECD countries”, OECD Social, Employment and Migration Working Papers, No. 141, OECD Publishing, Paris www.oecd.org/els/workingpapers.

Tolke, A. and M. Diewald (2003) “Insecurities in Employment and Occupational Careers and their Impact on the Transition to Fatherhood in Western Germany.” *Demographic Research* 12: 41-68.

Toulemon, L. (2011) “Should governments in Europe be more aggressive in pushing for gender equality to raise fertility? The first YES.” *Demographic Research* 24(7): 179-200.

Toulemon, L., A. Pailhé, C. Rossier (2008) “France: High and stable fertility”, *Demographic Research* 19(16): 503-556. Special Collection 7: Childbearing Trends and Policies in Europe.

Toulemon, L., and S. Pennec (2010) “Multi-residence in France and Australia: Why count them? What is at stake? Double counting and actual family situations.” *Demographic Research* 23,1:1-40.

van de Kaa, D. (2006) “Temporarily new: On low fertility and the prospect of pro-natal policies.” *Vienna Yearbook of Population Research* 193-211.

Table 1. Characteristics of Parity-zero Women Ages 18-34 and of Men Partnered to Parity-zero Women Ages 18-34, Weighted

	United States			Australia			Germany		
	Unpartnered women	**Partnered Women**	**Partnered Men**	**Unpartnered women**	**Partnered Women**	**Partnered Men**	**Unpartnered women**	**Partnered Women**	**Partnered Men**
Mean age**	25.0	27.6	30.3	24.2	27.6	31.1	25.6	29.3	34.2
Median age**	25	28	30	23	28	30	25	30	33
Education**									
Primary									
Secondary	47.9	37.3	42.6	65.8	48.6	58.8	43.6	43.1	47.7
Tertiary	-	-	-	34.2	51.4	41.2	56.4	56.9	52.3
Tertiary, < Bachelors	33.3	36.3	36.4	-	-	-			
Tertiary, Bachelors+	18.8	26.4	21.0	-	-	-			
Annual Work Hours**									
<750	22.8	12.1	4.9	32.3	14.1	6.3	52.9	24.3	7.5
750-1,749	24.7	21.5	10.0	25.4	17.5	9.0	12.7	18.8	9.0
1,750+	52.4	66.5	85.1	42.3	68.4	84.7	34.4	56.8	83.4
First births (percentage of t-1,t year pairs)	5.2	18.6	18.5	1.4	13.7	15.5	3.7	21.6	20.7
Unweighted N*	2,145	726	688	4,696	3,791	3,155	7,830	1,167	1,025
Percentage of women unpartnered and partnered***	74.7	25.3		59.9	40.1		87.6	12.4	

Table 1 continued

	Switzerland			Norway			France	
	Unpartnered women	**Partnered Women**	**Partnered Men**	**Unpartnered women**	**Partnered Women**	**Partnered Men**	**Unpartnered women**	**Partnered Women**
Mean age**	24.4	28.8	34.2	23.4	25.6	29.2	26.3	29.7
Median age**	24	29	33	22	25	28	23	27
Education**								
Primary				24.6	16.0	17.9	22.0	15.1
Secondary	81.5	67.3	52.7	45.4	37.1	45.4	49.3	40.7
Tertiary	18.5	32.7	47.3	30.0	46.9	36.7	28.7	44.2
Tertiary, < Bachelors				-	-	-	-	-
Tertiary, Bachelors+				-	-	-	-	-
Annual Work Hours**								
<750	28.4	8.5	3.5	56.4	33.4	17.3	62.8	35.4
750-1,749	10.1	15.2	5.4	15.6	20.2	12.7	10.6	16.3
1,750+	61.5	76.3	91.1	28.0	46.4	69.9	26.7	48.4
First births (percentage of t-1,t year pairs)	0.4	11.8	15.6	0.7	9.6	8.5	1.1	12.6
Unweighted N*	1,737	919	637	2,640	1,647	1,925	5,816	2,707
Percentage of women unpartnered and partnered***	63.4	36.6		61.6	38.4		68.2	31.8

Notes:

* These sample numbers (Ns) are of year-pairs of risk of first birth. They match the Ns in the regression models.

** All variables are measured in year t-1 of a year (t-1,t) pair of observations.

*** Weighted % partnered and unpartnered except for Norway and France, for which %s are unweighted.

Table 2a Logisitic Regression Parameter Estimates for the Annual Risk of First Birth, Unpartnered Women

	United States		Australia		Germany		France	
	coeff.	p-value	coeff.	p-value	coeff.	p-value	coeff.	p-value
Intercept	0.034	0.963	-3.856 **	0.000	-3.222 **	0.000	-4.394 **	<0.001
Age-25	-0.125 ***	0.000	0.004	0.885	0.118 **	0.000	0.027	0.456
(Age-25) Squared	0.008	0.169	-0.002	0.719	-0.006 †	0.083	-0.006 *	0.038
Educational attainment								
(reference: < Upper secondary)								
Secondary School							-0.516	0.102
Tertiary							-0.573	0.113
(reference: <= Upper secondary)								
Tertiary			-1.259 **	0.001	-0.006	0.691		
(reference: <= High school grad.)								
Some College	0.857	0.077						
Bachelors degree and above	-1.140 *	0.017						
Work hours in the previous year								
< 750 hours	-0.038	0.892	0.575 *	0.032	-0.280 †	0.090	-0.826 *	0.023
part-time/year hours	0.293	0.246	-0.111	0.733	0.080	0.643	0.912	
(reference: full-time/year hours)								
N (year-pairs)	2,145		4,696		7,830		5,816	

Notes:
** p < .01, * p < .05, † p< 0.10

All regression estimates are weighted.

Table 2b Logisitic Regression Parameter Estimates for the Annual Risk of First Birth, Partnered Women

	United States		Australia		Germany		Switzerland		Norway		France	
	coeff.	p-value	coeff.	p-value	coeff.	p-value	coeff.	p-value	coeff.	p-value	coeff.	p-value
Intercept	-1.071	0.294	-2.095 **	<0.001	-0.822 **	<0.001	-2.454 **	<0.001	-2.333 ***	<,0001	-4.394 **	<0.001
Age-25	-0.015	0.714	0.037 *	0.027	-0.107 *	0.017	0.112 †	0.073	0.130 ***	<0.001	0.027 **	0.001
(Age-25) Squared	-0.011	0.104	0.008 **	0.004	0.007	0.209	0.000	0.996	-0.006	0.220	-0.006 **	<0.001
Educational attainment												
(reference: < Upper secondary)												
Secondary School									-0.578 *	0.029	-0.516	0.167
Tertiary									-0.476 †	0.064	-0.573	0.156
(reference: <= Upper secondary)												
Tertiary			-0.269 *	0.011	0.024	0.158	-0.078	0.719				
(reference: <= High school grad.)												
Some College	0.044	0.849										
Bachelors degree and above	0.506 *	0.042										
Work hours in the previous year												
< 750 hours	0.099	0.743	0.199	0.170	-0.470 *	0.022	0.358	0.380	-0.660 **	0.007	-0.295 *	0.031
part-time/year hours	0.025	0.919	0.407 **	0.001	-0.416 †	0.051	0.170	0.546	-0.200	0.387	-0.439	
reference: full-time/year hours												
N (year-pairs)	726		3,791		1,167		919		1,647		2,707	

Notes:
** p < .01, * p < .05, † p< 0.10

All regression estimates are weighted.

Table 2c Logisitic Regression Parameter Estimates for the Annual Risk of First Birth, Partnered Men

	United States		Australia		Germany		Switzerland		Norway	
	coeff.	p-value	coeff.	p-value	coeff.	p-value	coeff.	p-value	coeff.	p-value
Intercept	-1.857	0.116	-1.878 **	<.001	-0.637 **	0.002	-2.107 **	<.001	-3.678 ***	<.001
Age-25	0.004	0.925	0.064 **	<.001	-0.084 *	0.011	0.070	0.189	0.118 ***	0.001
(Age-25) Squared	-0.002	0.511	-0.003 **	0.005	0.002	0.158	-0.003	0.258	-0.005 *	0.014
Educational attainment										
(reference: < Upper secondary)										
Secondary School									0.505 †	0.075
Tertiary									0.526 †	0.069
(reference: <= Upper secondary)										
Tertiary			-0.201 †	0.062	0.020	0.233	0.379	0.379		
(reference: <= High school grad.)										
Some College	0.411 †	0.084								
Bachelors degree and above	0.752 **	0.004								
Work hours in the previous year										
< 750 hours	-0.083	0.858	0.134	0.511	-0.694 *	0.018	-1.175	0.256	-0.608 *	0.047
part-time/year hours	-0.361	0.315	0.279 †	0.094	-0.086	0.763	-0.281	0.571	-0.061	0.827
reference: full-time/year hours										
N (year-pairs)	688		3,155		1,025		637		795	

Notes:
** p < .01, * p < .05, † p< 0.10

All regression estimates are weighted.

Table 1. Characteristics of Parity-zero Women Ages 18-34 and of Men Partnered to Parity-zero Women Ages 18-34, Weighted

	United States			Australia			Germany		
	Unpartnered women	**Partnered Women**	**Partnered Men**	**Unpartnered women**	**Partnered Women**	**Partnered Men**	**Unpartnered women**	**Partnered Women**	**Partnered Men**
Mean age**	25.0	27.6	30.3	24.2	27.6	31.1	25.6	29.3	34.2
Median age**	25	28	30	23	28	30	25	30	33
Education**									
Primary									
Secondary	47.9	37.3	42.6	65.8	48.6	58.8	43.6	43.1	47.7
Tertiary	-	-	-	34.2	51.4	41.2	56.4	56.9	52.3
Tertiary, < Bachelors	33.3	36.3	36.4	-	-	-			
Tertiary, Bachelors+	18.8	26.4	21.0	-	-	-			
Annual Work Hours**									
<750	22.8	12.1	4.9	32.3	14.1	6.3	52.9	24.3	7.5
750-1,749	24.7	21.5	10.0	25.4	17.5	9.0	12.7	18.8	9.0
1,750+	52.4	66.5	85.1	42.3	68.4	84.7	34.4	56.8	83.4
First births (percentage of t-1,t year pairs)	5.2	18.6	18.5	1.4	13.7	15.5	3.7	21.6	20.7
Unweighted N*	2,145	726	688	4,696	3,791	3,155	7,830	1,167	1,025
Percentage of women unpartnered and partnered***	74.7	25.3		59.9	40.1		87.6	12.4	

Table 1 continued

	Switzerland			Norway			France	
	Unpartnered women	**Partnered Women**	**Partnered Men**	**Unpartnered women**	**Partnered Women**	**Partnered Men**	**Unpartnered women**	**Partnered Women**
Mean age**	24.4	28.8	34.2	23.4	25.6	29.2	26.3	29.7
Median age**	24	29	33	22	25	28	23	27
Education**								
Primary				24.6	16.0	17.9	22.0	15.1
Secondary	81.5	67.3	52.7	45.4	37.1	45.4	49.3	40.7
Tertiary	18.5	32.7	47.3	30.0	46.9	36.7	28.7	44.2
Tertiary, < Bachelors				-	-	-	-	-
Tertiary, Bachelors+				-	-	-	-	-
Annual Work Hours**								
<750	28.4	8.5	3.5	56.4	33.4	17.3	62.8	35.4
750-1,749	10.1	15.2	5.4	15.6	20.2	12.7	10.6	16.3
1,750+	61.5	76.3	91.1	28.0	46.4	69.9	26.7	48.4
First births (percentage of t-1,t year pairs)	0.4	11.8	15.6	0.7	9.6	8.5	1.1	12.6
Unweighted N*	1,737	919	637	2,640	1,647	1,925	5,816	2,707
Percentage of women unpartnered and partnered***	63.4	36.6		61.6	38.4		68.2	31.8

Notes:

* These sample numbers (Ns) are of year-pairs of risk of first birth. They match the Ns in the regression models.

** All variables are measured in year t-1 of a year (t-1,t) pair of observations.

*** Weighted % partnered and unpartnered except for Norway and France, for which %s are unweighted.

Table 2a Logisitic Regression Parameter Estimates for the Annual Risk of First Birth, Unpartnered Women

	United States		Australia		Germany		France	
	coeff.	p-value	coeff.	p-value	coeff.	p-value	coeff.	p-value
Intercept	0.034	0.963	-3.856 **	0.000	-3.222 **	0.000	-4.394 **	<0.001
Age-25	-0.125 ***	0.000	0.004	0.885	0.118 **	0.000	0.027	0.456
(Age-25) Squared	0.008	0.169	-0.002	0.719	-0.006 †	0.083	-0.006 *	0.038
Educational attainment								
(reference: < Upper secondary)								
Secondary School							-0.516	0.102
Tertiary							-0.573	0.113
(reference: <= Upper secondary)								
Tertiary			-1.259 **	0.001	-0.006	0.691		
(reference: <= High school grad.)								
Some College	0.857	0.077						
Bachelors degree and above	-1.140 *	0.017						
Work hours in the previous year								
< 750 hours	-0.038	0.892	0.575 *	0.032	-0.280 †	0.090	-0.826 *	0.023
part-time/year hours	0.293	0.246	-0.111	0.733	0.080	0.643	0.912	
(reference: full-time/year hours)								
N (year-pairs)	2,145		4,696		7,830		5,816	

Notes:
** p < .01, * p < .05, † p< 0.10

All regression estimates are weighted.

Table 2b Logisitic Regression Parameter Estimates for the Annual Risk of First Birth, Partnered Women

	United States		Australia		Germany		Switzerland		Norway		France	
	coeff.	p-value	coeff.	p-value	coeff.	p-value	coeff.	p-value	coeff.	p-value	coeff.	p-value
Intercept	-1.071	0.294	-2.095 **	<0.001	-0.822 **	<0.001	-2.454 **	<0.001	-2.333 ***	<,0001	-1.555 **	<0.001
Age-25	-0.015	0.714	0.037 *	0.027	-0.107 *	0.017	0.112 †	0.073	0.130 ***	<0.001	0.068 **	0.001
(Age-25) Squared	-0.011	0.104	0.008 **	0.004	0.007	0.209	0.000	0.996	-0.006	0.220	-0.010 **	<0.001
Educational attainment												
(reference: < Upper secondary)												
Secondary School									-0.578 *	0.029	-0.263	0.167
Tertiary									-0.476 †	0.064	-0.269	0.156
(reference: <= Upper secondary)												
Tertiary			-0.269 *	0.011	0.024	0.158	-0.078	0.719				
(reference: <= High school grad.)												
Some College	0.044	0.849										
Bachelors degree and above	0.506 *	0.042										
Work hours in the previous year												
< 750 hours	0.099	0.743	0.199	0.170	-0.470 *	0.022	0.358	0.380	-0.660 **	0.007	-0.295 *	0.031
part-time/year hours	0.025	0.919	0.407 **	0.001	-0.416 †	0.051	0.170	0.546	-0.200	0.387	-0.439	
reference: full-time/year hours												
N (year-pairs)	726		3,791		1,167		919		1,647		2,707	

Notes:
** p < .01, * p < .05, † p< 0.10

All regression estimates are weighted.

Table 2c Logisitic Regression Parameter Estimates for the Annual Risk of First Birth, Partnered Men

	United States		Australia		Germany		Switzerland		Norway	
	coeff.	p-value	coeff.	p-value	coeff.	p-value	coeff.	p-value	coeff.	p-value
Intercept	-1.857	0.116	-1.878 **	<.001	-0.637 **	0.002	-2.107 **	<.001	-3.678 ***	<.001
Age-25	0.004	0.925	0.064 **	<.001	-0.084 *	0.011	0.070	0.189	0.118 ***	0.001
(Age-25) Squared	-0.002	0.511	-0.003 **	0.005	0.002	0.158	-0.003	0.258	-0.005 *	0.014
Educational attainment										
(reference: < Upper secondary)										
Secondary School									0.505 †	0.075
Tertiary									0.526 †	0.069
(reference: <= Upper secondary)										
Tertiary			-0.201 †	0.062	0.020	0.233	0.379	0.379		
(reference: <= High school grad.)										
Some College	0.411 †	0.084								
Bachelors degree and above	0.752 **	0.004								
Work hours in the previous year										
< 750 hours	-0.083	0.858	0.134	0.511	-0.694 *	0.018	-1.175	0.256	-0.608 *	0.047
part-time/year hours	-0.361	0.315	0.279 †	0.094	-0.086	0.763	-0.281	0.571	-0.061	0.827
reference: full-time/year hours										
N (year-pairs)	688		3,155		1,025		637		795	

Notes:
** p < .01, * p < .05, † p< 0.10

All regression estimates are weighted.