Ramifications of Divorce on the Economic Activity of Men and Women - A Multilevel Analysis

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Abstract

By employing linear multilevel analysis for repeated measures with growth models, logistic multilevel analyses, and fixed-effects models, this study analyzes how divorce affects different aspects of economic activity of men and women. Our multilevel analyses suggest that men's monthly salary and employment stability are hurt more following divorce, in comparison to women's. Nonetheless, our results are in line with previous research on the negative effect of divorce on women's economic status. This is because our fixed-effects models show, controlling for unobserved heterogeneity, that although women increase their employment stability and number of jobs held following divorce, in comparison to the times when they were married, their salary does not increase following marital disruption. It is also found that women experience a reduction in the growth rate of their salary following divorce. As for men, our fixed-effects models suggest that their employment stability is hurt following divorce, in comparison to their situation when they were still married, but that there are no substantial differences in men's salary following marital disruption. We also find that divorce does not affect the growth rate of the salary among men.

Ramifications of Divorce on the Economic Activity of Men and Women - A Multilevel Analysis

The discrepancies between the economic outcomes experienced by women and those experienced by men following divorce are well documented in previous studies which analyze the economic cost of marital disruption for men and women or for husbands and wives (Bartfeld, 2000; Burkhauser et al., 1990; Duncan & Hoffman, 1985; Morrison & Ritualo, 2000; Rowe, 1991; Smock, 1994). These studies mainly analyze how marital dissolution affects the economic well-being of the respondents, and do this mainly by comparing the own income or per-capita income before and following divorce. The findings regarding the implications of divorce for women's economic well-being are consistent, and demonstrate that women are economically hurt by divorce. The findings regarding the implications of divorce for men are less conclusive (for review, see Bianchi et al., 1999). While McManus & DiPrete (2001) find that most men who separate do not experience gains in their standards of living, other studies find that the economic well-being of men improves following disruption (Burkhausr et al., 1990; Duncan & Hoffman, 1985; Peterson, 1996).

In order to better understand the financial aftermath of marital disruption, it is crucial to explore the mechanisms which contribute to the change in economic status. The current study proposes that men's and women's outcomes following divorce depend not only on the loss of partner's income, on the loss of incomes which are due to tax and transfer programs, or on the size of mandatory or voluntary support payments given to the ex-partner; but are also dependent on changes in their employment and economic activity. Although previous studies addressed the issue of labor force activity of single mothers (Gonzáles, 2004; Jenkins, 1992; Meyer & Rosenbaum, 2001; Shemesh, 2005), it was usually done by using cross-sectional surveys in which the current labor force characteristics of the single-mothers were analyzed and compared to the characteristics of married mothers or of single women with no children. These studies do not analyze how the economic activity of men and women following divorce changes in comparison to their economic activity prior to marital disruption.

In this study we employ a unique longitudinal database, which links between a 20 percent sample of the Israeli 1995 census, annual register-based data from the National Insurance Institute of Israel (NIII) and the Tax Authorities, and the registration of divorces from the Ministry of Religious Services and from the formal population registry of Israel. Our sample contains 61,709 Jewish Israeli men and women who married for the first time by age 45, in Israel, since 1987 and are currently married or divorced in the 1995 census.

The principal innovation in our approach is the implementation of Multilevel Analysis in the context of person-years (Level-1) which are nested within individuals (Level-2). For the first time in this context of study, we employ growth models and Multilevel analyses, to analyze the effect of divorce on the economic activity of men and women, i.e. their monthly salary, employment stability and number of jobs held. In addition to these Multilevel models we also employ fixed-effects models. More specifically, we aim at answering one main research question concerning the possible economic consequences of divorce:

1) Do the economic ramifications of divorce differ between men and women in terms of their economic activity?

And additional two sub-research questions:

2) Does divorce affect not only the average salary of individuals, but also the growth rate of their salary?
3) Is the effect of divorce on the growth rate different for men and women?

Research Review

One indirect way in which the empirical literature addresses the issue of the change in economic activity following divorce is by analyzing the labor supply of single-mothers (Flug & Kasir, 2006; González, 2004; Zussman & Frish, 2008). According to the Israeli Central Bureau of Statistics (2009), in 2008, single-mothers constituted 90 percent of all lone-parents families with children aged 0-24, and in the head of 58% of the lone families stood a divorced parent, in comparison to 13%, 19% and 10% of the single-parent families in which the head of the household is separated, widowed or single, respectively. Therefore, single-mother families constitute most of the lone-parents families, and are usually formed due to divorce.

According to González (2004), Western countries differ in the extent to which single mothers participate in the labor force, so that in some countries single mothers are much more likely to work than other women, while in others, single mothers are much less likely to work. For example, González finds that between the mid-1980s and the mid-1990s, in France, Austria and Luxemburg employment rates (ER) of single-mothers are high (>65%), and are also higher than the ER of married mothers and of women without children. In the US and in Israel ER of single-mothers are high, but are similar to the ER of the other groups of women, while in the Netherlands, the UK and Australia ER of single-mothers are very low, and are much lower than the ER of the other groups. Okun & Oliver (2009) suggest that on the one hand divorced women are supposed to have higher economic incentives to participate in the labor force, due to the negative economic consequences of divorce and the decline in their economic well being. Nonetheless, on the other hand, divorced women are expected to have lower rates of full-time employment, due to the higher difficulty they experience in coordinating between labor (especially if it is full-time) and non-labor work.

One possible explanation to the way in which divorce might affect the economic activity of lone parents is policy regulations aimed toward this group. One example is the American 'workfare' policies and the European 'activation policies'. These policies stem from the growing recognition that the desirability and usefulness of the traditional welfare payments, which were meant to protect underprivileged groups (among them lone parents) from falling into poverty, are questionable. As the number of recipients increased, and so is the dependence on the welfare state, the criticism toward these welfare policies emerged (Knijn et al., 2007). The new mentioned policies aimed at making previously dependent people self-sufficient and independent, by giving them incentives to work or even compelling them to seek and take jobs. Therefore, unconditional benefits were replaced with requirements to work and the social protection model was replaced with commitment to employment activation (Knijn et al., 2007; Lødemel & Trickey, 2000). These policies took different forms and appeared in the US, in certain versions, since the 1970s, but proliferated there under the term 'workfare' policies since 1996 (Barbier & Ludwig-Mayerhofer, 2004; Wiseman, 2000). In Europe, activation policies were introduced in the mid/late 1990s, first in Denmark (in 1994, 1998), and then were extended to other countries, such as Sweden, Norway, the Netherlands, UK, and France (Barbier & Ludwig-Mayerhofer, 2004; Knijn et al., 2007). In countries where these kinds of policies exist and are aimed also toward lone parents, we can expect to see an increase in the economic activity of men and women following divorce. Meyer & Rosenbaum (2001) find, for example, that the change in the tax and transfer policy in the US, which was aimed at encouraging work among single mothers, accounts for over 60 percent of the 1984 to 1996 increase in the weekly and annual employment of single mothers relative to single women without children. Moreover, González (2004) shows that higher benefits for working single mothers encourage employment, while benefits if not working show insignificant effect. Nonetheless, according to Knijn et al. (2007), the effectiveness of these legislations is questionable. In the Netherlands, for example, 70 percent of the lone parents who started to work following the activation legislation were still partly dependent on social assistance. Therefore, these lone parents, although they are engaged in paid work, work mostly part-time and earn insufficient income to become fully independent from social assistance. In addition, Shemesh (2005) finds that in Israel, the 2002-2003 policy changes which were aimed toward increasing labor force participation rates of single-mothers, increased the number of women who joined the civilian labor force, but most of the single-mothers could not find a suitable job and remained unemployed.

Other factors which were found in previous literature to affect single-mothers' employment and economic activity are level of education, the presence of young children, age, child-care expenditures, expected earnings and income; so that older and more educated single mothers are more likely to work, while more and younger children reduce employment, as do higher child-care expenditures. Moreover, those with higher expected earnings are much more likely to work, while those with higher income from other sources are less likely to work (González, 2004; Jenkins, 1992; Shemesh, 2005).

The Israeli Context

Okun & Oliver (2009), who analyze the relationship between paid work and family structure among Jewish women in Israel, between the 1960s to the 1990s, find that employment rates of divorced women in the main working ages (25-45) are higher than those of married women across the years. Moreover, they find that the growth of their employment rates is slower than the growth rates of married women, so that in 1995 their employment rates are quite similar. Their cross-sectional comparisons show that employment rates of divorced women grew from 70 percent in 1961 to 75 percent in 1972, 77 percent in 1983 and 75 percent in 1995, while married mothers' employment rates were 28, 33, 58 and 76 percent, respectively.

Israel is characterized by a growing percent of single parent families, so that in 2005 their number was 3.8 times higher than in the 1980's. Most of the single parents in Israel are Jewish (80 percent), half of them are Israeli-born and almost 30 percent of them immigrated to Israel from the former USSR after 1990 (Lotan, 2007). In addition, as was found in the US (Duncan & Hoffman, 1985), most single parents in Israel earn low salaries, and single parent families are over represented among poor families and among families who rely on welfare transfers (Lotan, 2007; Flug & Kasir, 2006).

According to Flug & Kasir (2006) and Zussman & Frish (2008), two main changes were made during the early 1990s in the policy aimed toward single parents in Israel. In 1992, the Single Parents Law was introduced and granted divorced parents special treatment, especially under the Income Support Law. Under this law, divorced women were eligible to a generous welfare support, and some of the eligibility limitations by minimum income level were cancelled, therefore expanding the number of divorced women who were eligible for welfare support. In 1994 and 1995 additional legislative changes eased the terms of eligibility of this group. According to these authors, these legislative changes served to reduce the labor supply of single mothers at all levels: their labor force participation rates declined, as did the number of hours worked, while their share of part-time employment rose. Nonetheless, there was a delay in the appearance of these effects, which were apparent mostly from 1995 and until 2003. These effects were stronger among previously unemployed, young and less educated women. However, for married women the upward trend in the labor supply continued. Therefore, poverty rates among single mothers declined, but at the expense of increased dependence on the welfare system and reduced labor income (Flug & Kasir, 2006; Zussman & Frish, 2008).

Although Israel is characterized as having high levels of employment-supportive policies aimed toward mothers (Stier et al., 2001), only in 2002 and 2003 the legislation aimed toward single-parents has changed toward decreased dependence on welfare transfers and increased labor force participation. In 2002 the terms of eligibility of single-parent families for benefits were changed, so that the benefit and attendant assistance were reduced, and the income disregard was cut. In 2003, programs to provide incentives for single mothers to return to work and remain in employment by subsidizing employment were introduced. All these were intended to reverse the trend of declined employment and encourage single mothers to return to the labor market (Flug & Kasir, 2006; Shemesh, 2005).

In this study we analyze whether and how divorce affects men's and women's salary, employment stability and number of jobs, in comparison to their economic activity while they were still married. As was already described above, this micro-level analysis is conducted on data from a period in which there were no external positive incentives to work among single-parents, but rather, due to legislative changes, there were disincentives to employment. Nonetheless, there was a delay in the effect of these legislative changes (Flug & Kasir, 2006; Zussman & Frish, 2008);

so that only from 1995 and on did these changes affect the labor supply of single parents. Therefore, it is reasonable to refer to our research window as a period in which no essential changes in the external incentives or disincentive to work occurred; and indeed, a dummy variable for the period 1992-1995 and a dummy for the period 1993-1995 show no significant effect on the relationship between family status and the different outcome variables[1].

Data

This research is based on a unique longitudinal database, which links between a 20 percent sample of the Israeli 1995 census, annual register-based data from the National Insurance Institute of Israel (NIII), and the registration of divorces from the Ministry of Religious Services and from the formal population registry of Israel. The census data provide information on marital status in 1995, and exact year of first and last marriages. In addition, it contains data on current (highest) educational level and certificate. The data from the NIII follows people over 13 years, between the years 1983-1995, and contains data on annual salary and monthly employment status (i.e. whether the person work/does not work and how many employee jobs are being held each month). This information is based on Israel's tax authority reports on salaried employees. Therefore, salary and employment data from the NIII for a particular year are not available for people who were not salaried employees in that year. In addition, the NIII's data contain information on different social benefits a person received in a certain year, among them maternity allowance for women. The registration of divorces is available from 1985 until 2007. Due to problems in the quality of the data

[1] Since these dummy variables have no interaction effect with family status, and their exclusion does not change the results for the other independent variables, these variables are not included in the regression analysis.

on salary and employment in the years prior to 1987, and because data is available to us only until 1995, in this study we follow people from the year 1987 and until 1995.

Our sample contains 61,709 Jewish Israeli men and women (each group constitutes 50 percent of the sample) who married for the first time by age 45, in Israel, since 1987, and are currently married or divorced in the 1995 census (the census was conducted in November 1995). These men and women contribute a total of 309,788 person years. People are followed from the year of first marriage until the year 1995 or until the year of their second marriages, if those occurred before 1995[2]. Therefore, second marriages are not included in this analysis. For the currently divorced people, only divorces which took place until 1994 (N=1,574 – 57% women, 43% men) are included in the analysis. All these men and women were salaried employees at least one month during their marriage years.

This analysis does not refer to separation because in Israel there is no legal status for separation other than divorce. Moreover, the fact that only salaried employees are included in the analysis is not expected to affect the generalization of the findings, because in Israel most employees are salaried employees (in 1995, for example, 84% of Jewish employed men and 90% of Jewish employed women were salaried employees (Central Bureau of Statistics, 1996)).

Variables

We analyze four different dependent variables; each represents a different aspect of economic activity: log of average monthly salary, employment stability, and number of jobs held.

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[2] Pre-marital cohabitation is not included in the analysis because in Israel it is apparent mostly among relatively young secular Jews; but only to a small extent (only 3.7 percent of all couples lived in cohabitation in 2003) and for a short period of time (Blush-Klienman & Sherlin, 1999; Fogel, 2005).

Log of average monthly salary – This variable is computed by taking the log of the ratio between the annual salary and the number of months in which the individual worked as salaried employee. The annual salary is real salary (in Israeli Shekels) which is computed from nominal salaries with the consumer price index of the year 2006.

Employment stability – This is a binary variable in which 1 represents working 12 months during the year (i.e. stable employment) and 0 represents working less than 12 months (i.e. employment instability). We use dichotomous rather than continuous form of this variable because in 53% of the person-years in our sample individuals worked 12 months, while for the remaining cases there was a uniform distribution along the different levels of employment instability (i.e. less than 12 months).

Number of jobs held – One way in which an individual can increase economic activity is by holding more jobs. This is a binary variable in which 1 represents holding 2 or more employee jobs during at least 2 months in a certain year, and 0 represents holding 2 or more employee jobs only in one month or not at all. (i.e. holding zero or only one employee job during at least 11 months in a certain year).

We employ several independent variables; some of them are used as control variables:

Lagged family status – This is the central independent variable, which describes the family status of the individual as measured at the end of the previous year. This variable receives the value 1 if the individual is divorced in the previous year and 0 if married.

Gender – This variable receives the value 1 for men and 0 for women.

Academic degree – This variable receives the value 1 if the person holds academic degree and 0 otherwise.

Time – This variable measures the amount of time in years that had elapsed from the marriage year. This variable is included in the regression analysis as Z-scores variable, because we found high correlation between the coefficient of the original time variable and the intercept.

Age – This is a time varying covariate, which represents the age of the respondent at time *t*. It is included in the regression equation in both its linear and quadratic form, to allow for non-linear effect on the outcome variables. Both the linear and quadratic forms are included as Z-scores variables, because we found high correlations between the coefficients of these variables and the intercept.

Birth in year t – An increase in employment stability when a woman is divorced, in comparison to when she was married, might be due to possible births the women had during marriage (non-marital fertility rates in Israel are very low). Therefore, it is important to control for births among women. This variable is computed using the reports on maternity allowances which women received in a certain year[3]. It receives the value 1 if the woman received maternity allowance and 0 otherwise. Because this variable is available only for women, its effect will be tested in separate models for women only.

Number of children – This variable measures the total number of children a woman had in the 1995 census, i.e. at the end of the research window. Because this study follows people only until their second marriages or until the year 1995, if second marriages did not occur, this variable serves as a proxy to the number of children the woman had within her first marriage, since non-marital fertility in Israel is low.

[3] In Israel, only women who were employed at least 10 months out of the 14 months prior to termination of employment during a pregnancy which leads to a live birth are eligible to receive a maternity allowance.

Ethnic origin and generation – This is a control variable which measures the ethnic origin and the generation of the respondent in the country. The full description of its 6 categories, which are included in the analyses as dummy variables, is available in previous study of the author (2010).

Immigration status – This is also a control variable which is composed of a set of dummy variables as follows:

Native born Israeli – For subjects who were born in Israel.

Immigrated before 1990 – For subjects who were not born in Israel and immigrated to the country before 1990.

Immigrated after 1990 – For subjects who were not born in Israel and immigrated to the country starting 1990. This category serves as the reference category.

The two control variables of immigration status and of ethnic origin and generation are included in the analysis because previous research shows that new immigrants from the former USSR, who immigrated to Israel at the beginning of the 1990s, are characterized by higher divorce rates in comparison to Israeli-born Jews or to other immigrants (Dovrin, 2005), and so are Western Jews, in comparison to Oriental Jews (Kraus, 2002; Peres & Katz, 1991). Nonetheless, the results for these variables are beyond the scope of this paper, and therefore are not presented and analyzed[4].

Table 1 presents the mean or percentage, and the standard deviation of the dependent and independent variables in this analysis.

[Table 1]

[4] These results are available from the author upon request.

Empirical Approach

In our study we implement multilevel analyses, to analyze how divorce affects the economic activity of men and women. For the analysis of the salary outcome variable we use Hierarchical Linear Models with growth-curve models[5], and for our two binary dependent variables of employment stability and number of jobs we use Hierarchical Logistic Models[6]. The implementation of growth-curve models in a multilevel context is similar to the methodological approach implemented by Cherlin et al. (1998), in their study on the effects of parental divorce on mental health. Although some previous studies on the economic aftermath of divorce used fixed-effects methods (see, for example, Morrison and Ritualo, 2000), none of them used both fixed effects and a multilevel implementation, which also combines growth-curves models.

For our monthly salary dependent variable we use Linear Mixed-Models (LMM) for repeated measures using the SPSS software; and for our two binary employment stability and number of employee jobs variables we use logistic multilevel analysis using the MLwiN software. The intraclass correlation coefficients[7] of the different fully unconditional models of the dependent variables show that the variance between individuals constitutes between 70 to 77 percent of the variance in the dependent variables, and are significant. Therefore, there is a justification to conduct a multilevel analysis rather than just use a regular regression analysis. The decision to use LMM and not other widely used procedures, such as General Linear

[5] A thorough description of the Hierarchical Linear Models (HLM) method in general, and the implementation of growth-curve models in HLM in particular, can be found in Bryk & Raudenbush (1992).

[6] A review of this method and examples can be found in Guo & Zhao (2000).

[7] The intraclass correlation coefficient is computed by $\rho = \frac{\tau_{00}}{\tau_{00} + \sigma^2}$, i.e. as the ratio between the variance between subjects divided by the total variance of the dependent variable.

Models (GLM), stems from several reasons. First, the LMM does not assume independence of errors terms, which is a necessity in the case of repeated measures; second, while GLM assumes that all subjects are measured at the same points in time, LMM allows subjects to be measured at different points in time; third, LMM is asymptotically efficient for both balanced and unbalanced designs, but GLM is optimally efficient only for balanced designs.; fourth, LMM allows for a wide variety of assumptions about the covariance matrix; and, finally, it supports hierarchical data, while GLM does not (Garson, 2009).

The main reason for using 2-level hierarchical models stems from the recognition that the repeated measures of our outcome variables are nested within individuals (Bryk & Raudenbush, 1992:131). For our salary variable these models are also considered as linear growth models. At level 1, each person's outcome in a specific year depends on three parameters: family status in the previous year, the amount of time in years that had elapsed from the marriage year, and age

$$Salary_{ti} = \beta_{0i} + \beta_{1i}(Famstat)_{ti} + \beta_{2i}(Time)_{ti} + \beta_{3i}(Age)_{ti} + \beta_{4i}(Age^2)_{ti} + e_{ti} \qquad [1.1]$$

where β_{0i} is the outcome for person *i* on the first year of marriage (and age equals 0)[8], β_{1i} is the effect of person *i*'s family status in a certain year on his outcome in the same year, β_{2i} is the growth rate for person *i* over the data-collection period and represents the expected change during a fixed unit of time. This growth is assumed to be linear, due to the relatively short time frame of this study. Therefore, at level 1 the results for the time variable can also be interpreted as representing individual growth trajectories (Bryk & Raudenbush, 1992:134). The inclusion of the time variable is also meant to capture the secular time trend that the salary, for example, increase with time. The

[8] In our study, the time and age variables were included as Z-scores variables (see further description in the Variables section), therefore β_{0i} represents the outcome for person i with average time elapsed from marriage and average age.

coefficients for age represent the effect of age on the outcome variables, which is assumed to be non-linear. The term e_{ti} represents the specific error of individual *i* at time *t*. In the case of repeated measures of the salary variables, where we use linear models, we may suspect that error terms within an individual are correlated. Therefore, a reasonable choice of the residual error covariance will be a block diagonal matrix, where each block is a first-order autoregressive (AR1) covariance matrix (SPSS, 2002). Ignoring this clustering in the data and violating the independence assumption will result in underestimations of the standard errors (Guo & Zaho, 2000).

The unit of measurement in the level-1 model is not the individual but rather a person-year, or an observation on the individual at one point in time. The multilevel method allows us to estimate how characteristics of the individual (level-2 characteristics, such as gender, education, etc.) modify the values of β_{0i} and the values of the other β_{ni} coefficients (such as β_{1i}, β_{2i}, etc.). This is done in a level-2 model in which the unit of observation is the individual and the dependent variables are the β_{0i} and the β_{ni} parameters themselves. These parameters are treated as random effects and are allowed to vary between individuals. Because our main research question asks whether the effect of the family status (divorced vs. married) on the outcome variables is different for men and women, we will treat β_{1i} as a random effect which is allowed to vary at Level-2 as a function of gender. Therefore, the random-coefficient equation is

$$\beta_{1i} = \gamma_{10} + \gamma_{11}(Gender)_i + u_{1i} \qquad [1.2]$$

where γ_{10} represents the effect of divorce among women with average time elapsed from marriage; γ_{11} represents the differential effect of divorce on men in comparison to women; and u_{1i} represents the random effect. This interaction term does not appear in the models for women only, where the effect of divorce on women's salary is represented by the coefficient of the family status variable.

In addition, in order to test whether the growth rate of the salary changes when individuals are divorced in comparison to when they were married, we will also treat the coefficient of the time variable as a random effect, and let it change with family status (eq. 1.3). Moreover, in order to test whether the effect of the family status on the growth rate of the salary is different for men and women we will also include an interaction term with gender, and an interaction term of gender and family status in the time's slope equation

$$\beta_{2i} = \gamma_{20} + \gamma_{21}(Famstat)_{ti} + \gamma_{22}(Gender) + \gamma_{23}(Gender * Famstat)_{ti} + u_{2i} \qquad [1.3]$$

so that γ_{20} represents the growth rate of women when they are married, γ_{21} represents the differential growth rate of divorced women, in comparison to married women, γ_{22} represents the differential growth rate of married men in comparison to married women, γ_{23} represents the differential effect of divorce on the growth rate of men, in comparison to its effect on the growth rate of women, and u_{2i} is a random effect.

It is reasonable to assume that the intercept parameter β_{0i}, i.e. the average outcome of individual *i* while married, also varies as a function of Level-2, personal characteristics of individuals. Therefore, we will also treat β_{0i} as a random effect and let it vary as a function of the level-2 variables gender, education, ethnic origin and generation and immigration status

$$\begin{aligned}\beta_{0i} &= \gamma_{00} + \gamma_{01}(Gender) + \gamma_{02}(Academ) + \gamma_{03}(Ethnicity) \\ &+ \gamma_{04}(\mathrm{Im}\,migration) + u_{0i}\end{aligned} \qquad [1.4]$$

In equation 1.4, γ_{00} represents the grand mean and u_{0i} is the random effect.

In sum, we see that individual parameters become the outcome variables in a Level-2 model, where they may depend on some person-level characteristics. The full linear multilevel model is therefore

$$\begin{aligned}
&Salary_{ti} = \gamma_{00} + \gamma_{01}(Gender) + \gamma_{02}(Academ) + \gamma_{03}(Ethnicity) \\
&+\gamma_{04}(\mathrm{Im}\,migration) + \gamma_{10}(Famstat)_{ti} + \gamma_{20}(Time)_{ti} + \beta_{3i}(Age)_{ti} \\
&+\beta_{4i}(Age^2)_{ti} + \gamma_{11}(Famstat * Gender)_{ti} + \gamma_{21}(Time * Famstat)_{ti} \\
&+\gamma_{22}(Time * Gender) + \gamma_{23}(Time * Famstat * Gender)_{ti} + u_{0i} + u_{1i} + u_{2i} + e_{ti}
\end{aligned} \qquad [1.5]$$

In order to control for the possible effects of a birth in year t and of total number of children on women's outcomes, we will run separate models for women only, which include these two variables in their random intercept equations.

For the linear model of the salary, the covariance structure which yields the lowest Schwarz's Bayesian Criterion (BIC) for the random effects is the Scaled Identity. This structure has constant variance, and there is assumed to be no correlation between any elements. This is a common assumption when modeling the interaction of a random factor (such as year) with a fixed grouping factor (such as individual), where it is assumed that year*person interaction effect is normally distributed around a mean of zero, with unknown variance to be estimated (Garson, 2009). In addition, comparisons of the BICs values of different models show that a model in which both the intercept and the chosen slope parameters are allowed to vary fits the data better than a model that sets the intercept or slopes to the same value for all individuals.

As mentioned above, for our two binary response variables, employment stability and number of employee jobs, we use Hierarchical Logistic Models. The hierarchical models for these two response variables do not include a random effect of the growth term, therefore they include only a random intercept model and a random slope model for the family status variable, so that the logit models take the form

$$\log it(\pi_i) = \log\left(\frac{\pi_i}{1-\pi_i}\right) = \gamma_{00} + \gamma_{01}(Gender) + \gamma_{02}(Academ) + \gamma_{03}(Ethnicity)$$
$$+\gamma_{04}(\mathrm{Im}\,migration) + \gamma_{10}(Famstat)_{ti} + \beta_{2i}(Time)_{ti} + \beta_{3i}(Age)_{ti} \qquad [1.6]$$
$$+\beta_{4i}(Age^2)_{ti} + \gamma_{11}(Famstat * Gender)_{ti}$$

As in the linear model, we will show separate models for women only, which include in the random intercept models also the variables birth in year *t* and number of children in 1995.

It is important to note that because our three dependent variables might be endogenous, the regression models for each independent variable do not include the other dependent variables as explanatory variables. Separate analyses, which are not presented here, show that the exclusion of the remaining dependent variables from these regression models does not affect the results for the outcome variable under analysis.

Fixed-Effects Models

According to Allison (2009), in a random-effects model, the unobserved variables are assumed to be statistically independent of all the observed variables. Equations 1.2, 1.3 and 1.4 assume that the unmeasured characteristics of individual i, represented by the error terms u_{oi}, u_{1i} and u_{2i}, are not correlated with the measured characteristics, X_{1i}, …, X_{ni}. If this assumption is violated, the models will produce biased and inconsistent estimators (Halaby, 2004). To eliminate the threat of unobserved heterogeneity bias, we also estimate fixed-effects models that examine only within-individual variation in salary, employment stability and number of jobs held over time and control for all time-constant differences between individuals. In these models, we analyze how changes in the individual's family status between *t-1* and *t* are related to changes in our outcome variables between *t-1* and *t*. Thus, time-invariant unobserved heterogeneity between individuals is ruled out as a source of bias because we difference within the same individual across time, rather than differencing across individuals at the same time period. The concrete differences between our multilevel models and the fixed-effects models will be discussed in the results section of the fixed effects models.

One potential concern is that our measures of the family status are endogenous to economic activity, meaning that divorce is itself a product of economic activity, rather than vice versa. Our multilevel and fixed-effect models provide a solution to this problem because the *t-1* family status is measured prior to the outcome variables at time *t*.

For the fixed-effects analyses, we use *xtreg* and *xtlogit* commands in STATA, to separately estimate the main effects of our two time-variant variables, i.e. time elapsed from marriage and family status, on the three dependent variables of this study. Because time-invariant variables, such as gender, cannot be included in the fixed-effects models, the results of these models are presented in separate models for both sexes, for men only and for women only. The models for women also include the time-variant variable of birth in year *t*.

Results

Salary

a. What affects the average outcome of individuals?

The random intercept parameters (β_{0i}) in the model for both sexes and in the model for women only, in table 2, show the effects of the level-2 covariates on the intercept parameter – which is the expected log of monthly salary for individuals in general, and for women only, respectively, with mean time elapsed since marriage and mean age. Our results show that controlling for all the other independent variables, men's monthly salary is almost two times higher ($e^{.624} = 1.87$) than women's; and academic degree increases the log of the monthly salary by 23 percent ($e^{.202} = 1.224$) in comparison to non-academic degree. As for the level-1 covariates, it is found that the main effect of family status does not have a significant effect on the salary $(\gamma_{10} = .100)$[9], and so is the main effect of time $(\gamma_{20} = -.008)$[10]. Age has an inverse U-shape effect on the salary.

The intercept model for women only shows non-significant effects for the time and family status covariates, but a significant positive effect for academic degree. A birth in year *t* decreases the log of the monthly salary of women by 15 percent ($e^{-.157} = 0.85$) and each additional child decreases it by 9 percent ($e^{-.098} = 0.906$).

b. Does the effect of divorce differ for men and women?

The random family status slope model (for the β_{1i} parameter) in table 2 shows that the effect of divorce on women's log of monthly salary is positive but not significant ($\gamma_{10} = .100$). The differential effect of divorce on men's salary is negative and significant ($\gamma_{11} = -.356^{***}$) in comparison to women's, so that divorce decreases men's log of monthly salary by 30 percent ($e^{-.356} = 0.70$) in comparison to the effect of divorce on women's salary. In order to test whether men's salary is being hurt following divorce in comparison to their salary while they were married we ran equivalent model in which men are the reference category (not presented). In this

[9] The γ_{10} parameter is also the intercept parameter in the random family status slope's equation; therefore it has the meaning of the effect of divorce on women's salary, as is discussed below.
[10] The γ_{20} parameter is also the intercept parameter in the random growth rate slope's equation; therefore it has the meaning of the growth rate of the salary of married women.

[Table 2]

model $\gamma_{10} = -.256***$, therefore divorce reduce the log of men's monthly salary by 23 percent ($e^{-.256} = 0.774$). This negative effect of divorce on men's salary might represent a divorce penalty, or the flip side of the marriage premium. Divorced men cannot enjoy the benefits of marriage anymore, i.e. they do not have a partner who mainly specializes in home production, and who allows them to acquire more market-specific human capital and earn higher wages (Hersch & Stratton, 2000).
Following divorce they have to invest more time in home production and in raising their children, therefore their ability to invest in market production is being hurt, and so are their wages. The non-significant effect of divorce on women's salary demonstrates the ambivalent relationship between divorce and the economic activity of women, as was presented by Okun & Oliver (2009). On the one hand, women are expected to increase their market production following divorce and earn more money as they become the main breadwinners of their family; but on the other hand, the fact that they are also the main caregivers puts constraints on their ability to invest more in labor market production.

c. Does divorce affect growth in the salary?

The random growth rate slope parameter in table 2 (β_{2i}) show that while the growth rate of the salary of married women is negative but not significant ($\gamma_{20} = -.008$), divorce significantly decrease the growth rate of women's salary, in comparison to their growth rate when they were married ($\gamma_{21} = -.106*$), so that their growth following divorce is 10 percent lower ($e^{-.106} = .899$). This effect remains negative and significant after controlling for births and number and children among women. Moreover, the model shows that the growth rate of the salary is 12 percent higher for married men in comparison to married women ($e^{.117} = 1.124$), although it is found that there is no significant difference in the effect of divorce on men's and women's growth in the salary ($\gamma_{23} = .075$). A parallel model in which men are the omitted category (not presented) show that there is no significant effect of divorce on men's growth (i.e. γ_{21} in this model equals -0.30, but is not significant). Therefore, although we find a significant negative effect of divorce on the salary of men in the mean time elapsed from marriage (γ_{11}), i.e. when we only refer to one point in time, a longitudinal consideration of the growth in men's salary show that there is no significant difference between the growth rate of salary of men when they are divorced, in comparison to the times in which they were married. These results might suggest that men who are about to divorce fair worse economically in the pre-disruption period than men who eventually stay married.

Employment stability

a. What affects employment stability among individuals?

The random intercept parameter (β_{0i}) in the models of employment stability for both sexes and for women only, in table 3, show the effects of the level-2 covariates on the intercept parameter – which is the log odds of having stable employment rather than unstable employment, for individuals in general, and for women only, respectively, with mean time elapsed since marriage and mean age. The results show that controlling for all the other independent variables, men's odds of having stable employment are 2.5 times higher than women's ($e^{.907} = 2.5$), and the odds of having stable employment among individuals who hold academic degree are 9 percent higher in comparison to those who do not hold academic degree ($e^{.085} = 1.09$). The main effect of the lagged family status variable is found to be positive but not significant, whereas the effect of time is positive and significant. An increase in one standard deviation in the time elapsed since marriage increases the odds of stable employment by 3 percent ($e^{.032} = 1.03$). In addition, as in the salary, age has an inverse U-shape effect on the odds of having stable rather than unstable employment. The results of the random intercept parameter for women show that academic degree increase the odds of stable employment among women by 18.5 percent, in comparison to non-academic degree ($e^{.170} = 1.185$). Moreover, a birth in year *t* decreases women's odds of employment stability in this year by 21.5 percent, in comparison to years in which they did not give birth ($e^{-.242} = 0.785$). Each additional child also reduces women's odds of stable employment by 13 percent ($e^{-.137} = 0.872$).

b. Does the effect of divorce differ for men and women?

The random family status slope parameter (β_{1i}), for employment stability, in table 3, shows that divorce has a positive but non-significant effect on women's employment stability ($\gamma_{10} = .078$). Therefore, there is no significant difference in divorced women's employment stability, in comparison to married women. Nonetheless, the effect of divorce on men's employment stability, in comparison to women's employment stability, is negative and significant ($\gamma_{11} = -.888***$). Therefore, divorce decreases men's odds of stable employment by 59 percent ($e^{-.888} = 0.41$) more than it decreases women's odds of stable employment. In order to test whether men's employment stability is being hurt following divorce, in comparison to their employment stability while they were married, we ran equivalent model in which men

[Table 3]

are the reference category (not presented). In this model $\gamma_{10} = -.810^{***}$, therefore divorced men's employment stability is 55 percent lower ($e^{-.810} = 0.445$) in comparison to married men's employment stability. These differences in the effect of divorce for men and women also support our claim that there exists a negative divorce wage premium for men. One way in which men invest in market production is by holding continuous employment. Following divorce, due to the loss of the benefits of marriage, this ability is being hurt, and so is men's salary.

Number of jobs

a. What affects the number of jobs individuals hold?

The random intercepts parameter for the number of jobs variable in table 3 show that men's odds of increasing the number of jobs, i.e. of holding at least two employee jobs during at least two months along the year, are 26 percent lower than women's ($e^{-.303} = 0.74$). Because our database does not allow us to know whether the job held is full-time or part time job, it is possible that those who hold at least two employee jobs along the year are holding at least one part-time job. If this is the case, the gender differences we find might suggest that in Israel, as in other Western-developed countries, part-time jobs are held mostly by women (Stier & Lewin-Epstein, 2000). Our results further suggest that holding an academic degree increases the odds of holding at least two jobs by 79 percent ($e^{.583} = 1.79$), maybe due to the fact that there are some high-quality professions, such as of school teachers, social workers and nurses, which are characterized by part-time employment, are mostly occupied by women, and require academic degree (Kraus, 2002). The random intercept model for women also supports this assumption, because it is found that the odds of holding at least two jobs is 2 times higher among women with academic degree, in comparison to women with non-academic degree ($e^{.757} = 2.13$). Moreover, this model shows that the odds of holding at least 2 jobs in more than 2 months across the year are 5 times higher among women who give birth in a certain year, in comparison to the years in which they did not give birth ($e^{1.617} = 5.04$). The effect of the number of children was not found to be significant.

b. Does the effect of divorce differ for men and women?

The random slope parameter of family status (β_{1i}), for the number of jobs outcome variable, in table 3, shows that the effect of divorce on the number of jobs held is not significant among women ($\gamma_{10} = -.001$). Moreover, the differential effect of divorce on the number of jobs held by men in comparison to women was also found to be insignificant ($\gamma_{11} = .187$). A model in which the gender variable is recoded so that men are the reference category shows that γ_{10} equals .186 for men and is not significant, i.e. divorce does not affect the number of jobs men hold. Nonetheless, the coefficient of the family status independent variable in the intercept model for women only, is found to be positive and significant ($\gamma_{10} = .354^{***}$). This coefficient has the meaning of the effect of divorce on the number of jobs a woman holds, controlling for births, number and children and other covariates. Therefore, with mean time elapsed since marriage and mean age, and controlling for educational degree, births and number of children, divorce increases women's odds of holding at least 2 jobs by 42 percent ($e^{.354} = 1.42$).

Fixed-effects models

The fixed-effects models control for unmeasured characteristics, which might be correlated with both economic activity and divorce. For example, unmeasured personality attributes of the individual, such as instability, lack of responsibility, laziness, addiction, etc., might affect both the individual's economic activity and the risk of divorce. Charles & Stephens (2004), for instance, find that the non-economic suitability of the spouse is related both to his/her economic instability (for example, experiencing a dismissal) and to the precipitation of divorce. Using fixed-effects models, we can examine the effects of divorce controlling for such time-invariant unmeasured characteristics of the individual. In other words, in these models we rule out time-constant unobserved heterogeneity between individuals and measure only the change within individual. Therefore, the advantage of the multi-level analysis is in its ability to explicitly model the differential effect of divorce between men and women. The fixed-effects models rule out unobserved heterogeneity, which is not accounted for in the multilevel analysis.

The fixed-effects models presented in Table 4 show that the effect of divorce on men's salary is negative but not significant, and its effect on women's salary is positive but not significant. These results are in line with the multilevel results for women, where we find positive but non-significant effect for women ($\gamma_{10} = .100$). The negative effect of divorce on men's salary is also in line with the multilevel results for men, where in a model in which men are the omitted category, we find negative and significant effect for men ($\gamma_{10} = -.255^{***}$)[11]. Nonetheless, this effect is not found to be significant in the fixed-effects model. It is possible that the significant

[Table 4]

[11] This effect remains negative and significant when we exclude the growth as outcome parameter from the equation.

negative effect we found among men in the multilevel analysis is due to unobserved characteristics of men, which affect both their salary and their propensity to divorce.

The results of the fixed effects models for employment stability are consistent with the results of the multilevel analysis for men only. The effect of divorce on men's employment stability is significantly negative in both models ($\gamma_{10} = -.810^{***}, \beta_{FE} = -.556^{***}$), i.e. following divorce men's odds of having stable employment are reduced in comparison to their situation when they were married. As for women, the effect of divorce on their employment stability is positive and significant in the fixed-effects model but negative and insignificant in the multilevel model, controlling for births ($\gamma_{10} = -.073, \beta_{FE} = .455^{***}$).

The fixed-effects models' results for the effect of divorce on the number of jobs men and women hold are consistent with the multilevel results. Both models show positive but non-significant effect for men ($\gamma_{10} = .186, \beta_{FE} = .138$), and positive and significant effect for women ($\gamma_{10} = .354^{***}, \beta_{FE} = .328^{**}$), controlling also for births.

To sum, overall it seems that our results for the fixed-effects models are consistent with the results of the multilevel models. Only in the analysis of the effect of divorce on men's salary the results were significant in the multilevel model but insignificant in the fixed-effects level. In the case of employment stability among women the effect of divorce was found to be positive and significant in the fixed-effects model rather than in the Multilevel model. Therefore, these results increase our confidence that the effect of divorce on men's and women's economic activity is not simply a result of unmeasured time-invariant characteristics of the individual, but is due to divorce itself. Table 5 summarizes the results of the multilevel analyses and the fixed effects analyses for the effect of divorce on the different aspects of economic activity of men and women.

[Table 5]

Summary and Conclusions

One of the side effects of the sharp increase in divorce rates in most Western-developed countries, in the last few decades, is the increase in the number of single parent families, most of them are headed by women. Due to the fact that single-mother families are over represented among poor families and among families who receive welfare support, most of the previous research focus on the profound economic implications of divorce for women and children and on the discrepancies between the economic outcomes experienced by women and those experienced by men (Bartfeld, 2000).

In this study we aim at understanding one of the main mechanisms through which divorce affect the economic well being of men and women. We analyze the effect of divorce on the economic activity of men and women, i.e. the effect of divorce on the monthly salary, on the employment stability and on the number of jobs held, among the Jewish population in Israel, in a period in which there were no profound changes in the external incentives or disincentives to work. Therefore, we analyze the net effect of divorce on the labor supply of men and women following divorce, which is not affected by external incentives to participate in the labor force. This study has profound implications for policy aimed toward single-parent families, because it gives better understanding as to what are the mechanisms through which legislation can promote labor supply of men and women following divorce, since increased labor force participation among single parents might prevent their families from falling into poverty, and therefore also decrease their reliance on the welfare state (Knijn et al., 2007).

Using linear multilevel analyses with growth models and logistic multilevel analyses, together with fixed-effects models, our analysis shows that following divorce women increase their economic activity by means of having more continuous and stable employment and by means of increasing the number of jobs held, in comparison to the times in which they were married. Although we would expect the increased stability and the increase in number of jobs to yield higher earnings for women, we find that their monthly salary does not increase as a result of these changes[12], and that in the long run, the increase in their salary is slower in comparison to the increase they experienced when they were married. It is possible that the increase in the number of jobs represents holding part-time employments, an information which is not available in our database. If this is the case, and our findings actually demonstrate an increase in part-time employment, this might explain why we could not find differences in women's salary, as part-time employment is found to be related to wage penalty, occupational segregation and an exclusion from most lucrative jobs, in most Western countries (Stier & Mandel, 2009). These results also point to the conflicting role single-mothers experience as mothers and main breadwinners, because on the one hand they have to economically support their family by increased labor supply, but on the other hand, the fact that they are also the main caregivers puts constraints on their ability to invest more time and effort in labor market production (Okun & Oliver, 2009). Therefore, they are being forced to seek for part-time jobs because these kinds of jobs enable mothers to combine both paid and unpaid work (Stier & Mandel, 2009). For that reason, policies aimed at promoting

[12] Similar results were found for the annual salary.

the labor supply of single mothers most also include components of extensive social support, such as day cares and afternoon child care facilities, which will help the mothers extend their labor force activities while their children are taken care of.

As for men, we find a reduction in men's employment stability following divorce, in comparison to the times in which they were married. Although we found a significant negative effect of divorce on the salary of men in the mean time elapsed from marriage, i.e. when we only referred to one point in time, a longitudinal consideration of the growth in men's salary show that there is no significant difference between the growth rate of salary of men when they are divorced, in comparison to the times in which they were married. The results of the fixed-effects model for salary support these findings. These results might suggest that men who are about to divorce fair worse economically in the pre-disruption period than men who eventually stay married. Nonetheless, the results of random family status slope model in the multilevel analysis show that men's salary is being hurt more following divorce in comparison to women's salary. These results suggest that following divorce men have to invest more time in home production and in raising their children, therefore their ability to invest in market production is being hurt, and so are their wages. This might be considered as the flip side of the marriage premium, or as divorce penalty, because this reduction might be caused due to the loss of the benefits of marriage, i.e. divorced men do not have a partner who mainly specializes in home production, and who allows them to acquire more market-specific human capital and earn higher wages (Hersch & Stratton, 2000).

These results for men and women do not contradict previous findings as for the discrepancies in the economic well-being of men and women following divorce. Women's well-being might be reduced following divorce because although they invest more in labor supply, they do not experience an increase in their earnings, and the growth of their salary becomes slower. Men, who overall experience a reduction in the salary following divorce, in comparison to women, were not found to be hurt by means of salary in comparison to their situation when they were married.

To better understand the marriage premium and negative divorce premium, it might be interesting to analyze in a future study whether second marriages rehabilitate the growth rate of the salary among women, in comparison to the times when they first married and divorced, and how the other aspects of economic activity of men and women change following remarriages.

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Tables

Table 1. Variables' means and standard deviations, computed over person-years (N=309,788)

Variable	Mean / Percentage	SD
Log of average monthly salary (in NIS)	8.308	1.336
Employment stability (stable=1, not stable=0)	.640	.479
Number of jobs (2 or more=1, otherwise=0)	.183	.387
Family status (*t-1*) (divorced=1, married=0)	.014	.118
Gender (1=male, 0=female)	.492	.499
Academic degree (1=academic, 0=otherwise)	.213	.409
Time (in years)	2.638	2.188
Age (in years)	28.335	4.913
Birth in year *t* (for women) (1=yes, 0=no)	.179	.384
Number of children in 1995 (for women)	1.301	1.040

Note: For the dichotomous variables, the mean represents the percentage of the value which is coded 1.

For the time, age and number of children variables means and SDs were computed using the original variables.

Ethnic origin and generation in the country, and immigration status, are included in the analysis but not presented.

Table 2. Effects of Selected Independent Variables on the Intercept and the Slope Parameters of a Linear Multi-Level Model Predicting Monthly and Annual Salary.

	Dependent Variables (Outcome variable = Log of Monthly Salary)				
	Model for both sexes			Model for Women only	
Independent Variables	Intercept Parameter (β_{0i})	Family Status Slope Parameter (β_{1i})	Growth rate Slope Parameter (β_{2i})	Intercept Parameter (β_{0i})	Growth rate Slope Parameter (β_{2i})
Family status in *t-1* (1=divorced)	.100 (.052)	-	-.106* (.042)	-.011 (.056)	-.105* (.046)
Gender (1=male)	.624*** (.008)	-.356*** (.080)	.117*** (.008)	-	-
Gender * Family status	-	-	.075 (.067)	-	-
Time since marriage	-.008 (.006)	-	-	.002 (.007)	-
Age in year *t*	.951*** (.033)	-	-	1.018*** (.048)	-
Age^2	-.762*** (.032)	-	-	-.821*** (.049)	-
Academic degree (1=yes)	.202*** (.009)	-	-	.173*** (.014)	-
Birth in year *t* (1=yes)	-	-	-	-.157*** (.009)	-
Number of children in 1995	-	-	-	-.098*** (.006)	-
Constant	7.736***	.100	-.008	7.942***	.002
Explained Level-2 variance	30%			42%	

Note: Numbers in parentheses are standard errors; N (all sample) = 61,709, N (women) = 31,024.
Ethnic origin and generation in the country, and immigration status, are included in the analysis but not presented.
* p < .05, ** p < .01, *** p < .001

Table 3. Effects of Selected Independent Variables on the Intercept and the Slope Parameters of a Logistic Multi-Level Model Predicting Employment Stability and Number of Jobs.

	Dependent Variables (Outcome variable = Employment Stability)			Dependent Variables (Outcome variable = Number of jobs)		
	Model for both sexes		Women only	Model for both sexes		Women only
Independent Variables	Intercept Parameter (β_{0i})	Family Status Slope Parameter (β_{1i})	Intercept Parameter (β_{0i})	Intercept Parameter (β_{0i})	Family Status Slope Parameter (β_{1i})	Intercept Parameter (β_{0i})
Family status in *t-1* (1=divorced)	.078 (.064)	-	-.073 (.064)	-.001 (.067)	-	.354*** (.081)
Time since marriage	.032*** (.007)	-	.042*** (.010)	-.182*** (.009)	-	-.058*** (.013)
Age in year *t*	2.044*** (.061)	-	1.862*** (.081)	.578*** (.076)	-	.541*** (.105)
Age^2	-1.688*** (.059)	-	-1.535*** (.080)	-.600*** (.074)	-	-.540*** (.105)
Gender (1=male)	.907*** (.015)	-.888*** (.101)	-	-.303*** (.018)	.187 (.129)	-
Academic degree (1=yes)	.085*** (.018)	-	.170*** (.024)	.583*** (.020)	-	.757*** (.027)
Birth in year *t* (1=yes)	-	-	-.242*** (.016)	-	-	1.617*** (.019)
Number of children in 1995	-	-	-.137*** (.010)	-	-	-.011 (.012)
Constant	-.532***	.078	-.307***	-1.631***	-.001	-2.153***
Explained Level-2 variance	24%		23%	29%		25%

Note: Numbers in parentheses are standard errors; N (all sample) = 61,709, N (women) = 31,024.
Ethnic origin and generation in the country, and immigration status, are included in the analysis but not presented.
* $p < .05$, ** $p < .01$, *** $p < .001$

Table 4. Fixed-Effects Estimates of the Effect of Divorce and the Time Elapsed from Marriage on the Outcome Variables.

	Outcome Variable								
	Both sexes			**Men only**			**Women only**		
Independent Variables	**Monthly Salary**	**Employment Stability**	**Number of Jobs**	**Monthly Salary**	**Employment Stability**	**Number of Jobs**	**Monthly Salary**	**Employment Stability**	**Number of Jobs**
Time (in years)	.088*** (.001)	.156*** (.004)	-.118*** (.004)	.103*** (.002)	.144*** (.006)	-.207*** (.007)	.068*** (.002)	.158*** (.005)	-.026*** (.006)
Family Status in *t-1* (1=divorced)	-.029 (.034)	.062 (.079)	.145 (.094)	-.069 (.051)	-.556*** (.124)	.138 (.148)	.009 (.046)	.455*** (.107)	.328** (.129)
Birth in year *t* (1=yes)	-	-	-	-	-	-	-.232*** (.010)	-.633*** (.021)	2.084*** (.026)

Note: The (centered) time variable was included in its regular form, and not in its Z-score form.
Numbers in parentheses are standard errors.
* p < .05,   ** p < .01,   *** p < .001

Table 5. A Summary of the Results of the Effect of Divorce on Men's and Women's Economic Activity, from Multi-Level and Fixed-Effects Analyses.

	Women when divorced vs. when married		Men when divorced vs. when married		Men vs. Women
	Multi-Level Model	Fixed-Effects Model	Multi-Level Model	Fixed-Effects Model	Multi-Level Model
Log of monthly salary	$+^{N.S.}$	$+^{N.S.}$	$-^{***}$	$-^{N.S.}$	$-^{***}$
Growth rate of salary	$-^{*}$		$-^{N.S.}$		$-^{N.S.}$

Employment stability	$-^{N.S.}$	$+^{***}$	$-^{***}$	$-^{***}$	$-^{***}$
Number of jobs	$+^{***}$	$+^{**}$	$+^{N.S.}$	$+^{N.S.}$	$+^{N.S.}$

Note: (-) = Negative effect of divorce
(+) = Positive effect of divorce
N.S = Not-significant
* p < .05, ** p < .01, *** p < .001