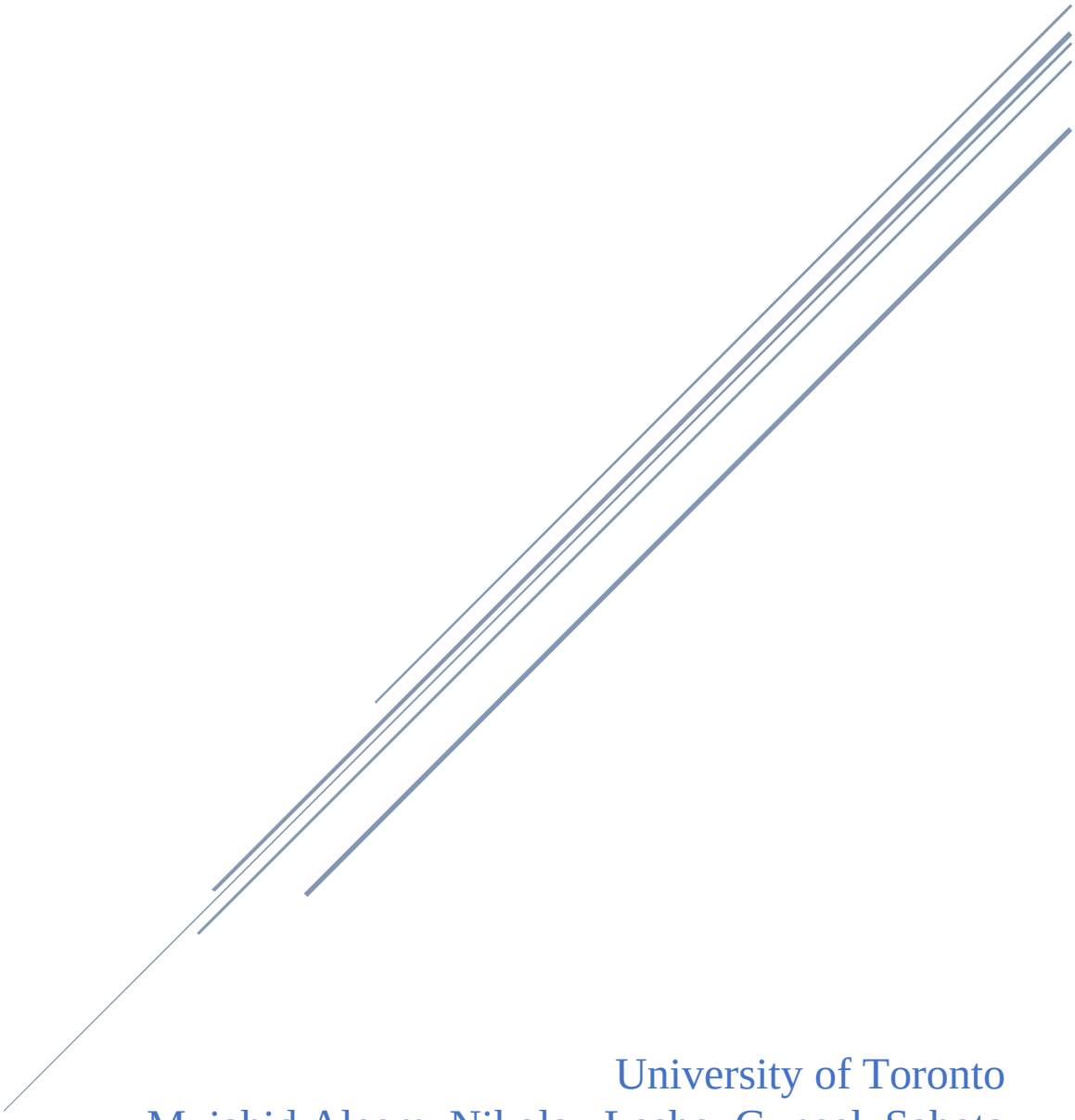
ANALYSIS OF ETHIOPIA'S CHILD MARRIAGE LAW

ECO403: Topics in Development Economics and Policy



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We would like to thank Professor Vivalt for her fundamental assistance towards this research project. This paper was made possible with her support and prepared for the ECO403 class.

Introduction

The practice of child marriage in poorer countries has been a prevalent issue in development economics. The perpetuation of child marriage has correlated with extreme poverty and has impeded efforts to achieve equity and economic growth. In McGavock's paper, "Here waits the bride? The effect of Ethiopia's Child marriage law" (2021), observes the causal impact of the "Revised Family Code" on the age of first marriage of women in Ethiopia. The marriage law was implemented by the Federal Democratic Republic of Ethiopia in July 2000, which raised the legal minimum age of marriage for women from 15 to 18 years old and left the men's minimum age of marriage unchanged at 18. The reform also guaranteed the equality of the spouses during the conclusion, duration, and dissolution of marriage. However, this law allowed for marriages of girls at 16 and 17 only with parental consent. In the paper, McGavock (2021) deduces the effect of Ethiopia's reform of child marriage, where she concluded that the law resulted in a decrease in the probability of women getting married earlier. This was attributed to delays in under-age marriages among the younger girls (i.e., marriages involving girls under age 16) which then led to lower total lifetime fertility.

To estimate the effect of the Revised Family Code, McGavock (2021) compares the average age of women married during the 10 years pre-and post-reform and then takes advantage of the staggered timing of the reform in each region (i.e., region derives from Ethiopia's nine semi-autonomous regions: Addis Abba, Dire Dawa, Oromia, Amhara, SNNPR, Benishangul-Gumuz, Tigray, Harari, and Gambela). In her report, McGavock (2021) primarily focuses on the short-term effect of the implementation of the reform as she only considers the period of 2000-2010 (i.e., post-reform period) and 1990-1999 (i.e., pre-reform period). We believe that to analyze the long-term effect of this law is to extend the post-reform period by approximately a decade. Therefore, we extend upon McGavock's (2021) study by including the year 2016 in the post-reform period. We then observe the average change in the first marriage age in women by comparing this to the pre-reform period (i.e., 1990-1999) to gain a sense of the long-term effects of the law.

The motivation for using more recent data on the reform is to determine if the long-term effects of the law match or differ from the findings that McGavock (2021) had obtained in her shorter-term analysis. In her report, she regarded the reform law as the main reason behind the delay in younger women's early marriages. By observing the long-term effects, we can account for changes in the educational attainment, fewer births, health investments, and employment of the married women surveyed in this study that may have not been observed in the short-term study. Therefore, the motivating factor behind our study is to analyze whether our results significantly differ from the results of McGavock's (2021) report and to also observe the long-term effects of the reform.

Lastly, our study provides a few contributions to the literature on child marriage. Our analysis revolves around the long-term effects of the law, which is an extension of the first causal analysis of the efficacy of an age of consent law in developing countries conducted by McGavock (2021) in the original paper. Therefore, the method that we observed the distributed roll-out of the reform across the semi-autonomous Ethiopian regions during 2000-2016 does provide the exogenous variation which is difficult to obtain from developing countries. Second, our study also provides a reasonable exogenous identification in the estimation process concerning the causal effect of early marriage among women on areas including fertility and education attainment. Hence, our paper contains the confirmation of causality as well.

Data

For our report, we retrieved the 2016 and 2000 household surveys from the Demographic and Health Surveys (DHS), and are conducted in alliance with the Central Statistical Agency in Ethiopia. These sources provided nationally and regionally representative cross-sections of women ages 15-49 and men ages 15-59 in 2000 and 2016 for our study. The women in these surveys were asked about their marital status and their "age at first cohabitation" — which the DHS then accounts in its records as age at first marriage for these "ever-married" women. Furthermore, the 2016 dataset (i.e., household surveys) includes respondents who account for the treatment group; and the 2000 dataset (i.e., household surveys)

includes respondents who account for the control group. Hence, the treatment group would be the respondents who were affected by the reform law (i.e., 2000-2016), while the control group would be the respondents who were not affected by the reform law (i.e., 1990-1999).

Methods

Our paper aims to utilize a difference-in-differences regression approach to represent the long-term effects of the marriage reform law introduced in July 2000. To achieve this, we observe the change in the age at first marriage of Ethiopian women and then regress this against the change in the proportion of Ethiopian women before and after the reform. In essence, as the proportion of Ethiopian women increases, we expect the age at first marriage to increase.

In the original paper, the author used Post-Reform as the dummy variable that indicated if the marriage took place before or after the reform — we do the same in our study. Essentially, we use the same regression equation as in the original paper to account for the fixed effects and the treatment effects. However, in our analysis, we use a different point in time (i.e., 2016) to compare the post-reform against the pre-reform period (i.e., 2000 that encompasses respondents from 1990-1999). Therefore, we have the "Age of First Marriage" variable as the dependent variable, the "Post-reform" variable (i.e., the indicator or dummy variable), and the constant term in our regression that encapsulates the fixed effects including the year of marriage (ηt) and year of survey ($\nu\tau$). Additionally, the β coefficient on the Post-reform dummy captures the treatment effect of the reform (i.e., change in the median age at first marriage due to the reform law). We used the most recent dataset in 2016 and the dataset published before the reform in 2000. Thus, we treat the respondents in 2016 as the treatment group and respondents in 2000 as the control group, as previously mentioned. Also, this is suitable as the reform was implemented at the same time as when the group was treated. Moreover, the study conducted in 2000 contains information from before the reform (i.e., 1990-1999), while the respondents in 2016 are all subjects who have experienced the reform.

Lastly, our treatment group will be the proportion of people when the Post-reform variable is active (i.e., the dummy variable is a binary variable that takes on "1" if active; "0" if not active). On the other hand, the control group is the remainder of the proportion of people given the case the Post-reform variable is equal to "0." Then, the first effects represent the age at first marriage if there was no reform or if the number of women surveyed was "0." Thus, this is considered as the minimum age of first marriage in our regression. Finally, we produce our difference-in-differences estimator which gives us the effects of the reform on the age at first marriage. Also, it is to be noted that in our study, we utilize this difference-in-differences estimator to provide us with the treatment effects (i.e., the effect of reform) which is parallel to the approach that McGavock (2021) uses in her regression with the β coefficient capturing the treatment effect. This estimator provides us with the average change in the median age at first marriage due to the reform law. In particular, the estimator is calculated by multiplying the Post-reform dummy variable by the proportion of respondents. Thus, this results in our regression equation where there is a dependent variable (Age at first marriage) and the predictor variables consisting of the Post-reform dummy variable and the proportion of respondents' variable along with the constant term which accounts for the fixed effects which were discussed earlier.

Results

Several trends were observed in the age of first marriage for both younger and older women between 2000-2016, outlined in the graphs in Figure 1. Among women who were 20-49 in the survey, the average age of marriage increased from 16.4 to 17.5 from 2000 to 2016. A similar increase appeared for women who were 25-49 in the survey, from 16 in 2000 to 17.1 in 2006. This is important as it highlights older individuals who may not have been married after the reform. In addition to an increase in the median age of marriage, a downward trend was observed in the percentage of women married by ages 15 and 18. This downward trend has been particularly pronounced for the younger generations, with the percentage of respondents aged 15-19 who were married by 15 dropped from 14.4% in 2000 to 5.7% in

2016. On the other hand, the percentage of respondents aged 15-19 who were married by 18 dropped from 49.1% in 2000 to 40.3% in 2016.

The results of the regression (Figure 2) indicate an increase in the age of marriage associated with the law, as the Post_Reform dummy variable coefficient, Beta, is positive. This indicates that the age of marriage increases by .2445827 on average when the law is active. This is consistent with the trend in the age of marriage observed from 2000 to 2016, with an increase in the average age of marriage and a decrease in the proportion of those married by age 15 among the younger generation. The value of the Beta coefficient is statistically significant, as indicated by the P > |t| section, where it is 0.821, which is greater than 0. The coefficient for the variable representing the proportion of respondents, Gamma, is also positive, indicating that as the proportion of people increases, the age of marriage also increases, at .112. However, this value is statistically insignificant, with a P > |t| value of 0. Overall, there was a statistically significant change in the age of marriage following the reform, as the estimator variable had a coefficient of 0.0116783 with a P > |t| value of 0.449. Therefore, the long-term effect of the reform results in a continued increase in the age of marriage – but at a more gradual rate as compared to the sudden increase that McGavock observed in her findings.

Discussion & Limitations

The analysis has been extended from the original paper by using data from the DHS website for the years 2016 and 2019. However, although the data is fully available for the year 2016, the processing analysis of the data collected is still ongoing for the year 2019. Furthermore, while the fieldwork to collect the data was conducted from January to June in 2016, it was only conducted from the months of March to June in 2016, adding further limitation to the availability of data from 2016.

There are various limitations with respect to our extension that builds on McGavock's (2021) report. As we utilize a regression like the author's with the exception of some missing information of the regional effects of the reform, the original paper's limitations are also applied to our report. Thus, there

are two major assumptions which are required for the methods to deliver a causal estimate: 1) the reform's timing was exogenous to regional trends and 2) other policies and interventions do not explain the results of delayed early marriages observed. Furthermore, the author assumes exogeneity to regional trends (i.e., regional fixed effects), this enables us to provide empirical evidence as the timing of the reform is not observed to be systematically related to pre-existing norms or trend rates in early marriage. However, as we do not include region-specific time trends (i.e., regional fixed effects) due to some missing information, our analysis excludes the possibility of the correlation between the timing of the reform and the pre-existing norms/trend rates in early marriages. Therefore, there could be possible limitations to the extent that we consider the effect of the reform law as there could be dynamic treatment effects present. Additionally, there have been various government and NGO interventions in Ethiopia around the time frame when the family law reform was implemented - thus, resulting in some impact on child marriage. However, to enable for the minimum impact of these interventions, like the author, we did not include data from locations in Ethiopia consisting of such interventions. So, this could result in some limitation regarding our causal estimate as this is a strong assumption which is made. Lastly, as we do not incorporate the regional fixed effects in our regression, factors affecting marriage age such as economic conditions, changes in norms over time, and poverty levels which are specific to regions are not flexibly controlled by the regional fixed effects. Thus, this poses another limitation regarding our analysis as it does not examine the impact of the law to the greatest extent as it does not consider the stated factors.

The results from our report propose that alongside the legal reforms (i.e., Revised Family Code), one can increase the effectiveness of the law by combining it with legal literacy operations which could improve upon behavior towards early marriage and outcomes against the strong cultural practices and poor administration of regions within developing countries. So, introducing some form of an information campaign (i.e., NGO efforts directed towards educational attainment) after the implementation of the change in legal age of consent/age at first marriage can assist in sustaining the positive long-term effects of such reform policies. Moreover, the aspect of increasing the legal minimum age of marriage of the

reform law can be utilized as an effective family reform law in other developing countries where rates of early marriages is extremely high (e.g., India).

Conclusion

In all, the long-term effects of the reform show that there was a decrease in the age of first marriage for women on average. The reform disproportionately affected the youngest Ethiopians as well. This is in line with the conclusions made by McGavock, wherein her report she attributed to having a decrease in the age of first marriage as well. Like the original paper, however, our regression only attributes the change in the age of marriage due to the reform, which could accentuate an omitted variable bias considering the prolonged observational period. Omitted factors such as changes in political landscapes, economic prospects, and changes in culture carry more significance when observing the effects of a policy long term. In the future, when reviewing the causal impact of a policy such as the Revised Family Code, introducing an IV design to reduce omitted variable biases could be insightful.

References

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Figures

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. reg AgeofFirstMarriage Post_Reform proportionOfRespondents estimator
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Source	SS	df	MS		
Model	553.220754	3	184.406918	Number of obs	= 58
Residual	124.848212	54	2.31200392	F(3, 54)	= 79.76
				Prob > F	= 0.0000
				R-squared	= 0.8159
				Adj R-squared	= 0.8056
Total	678.068966	57	11.8959468	Root MSE	= 1.5205

AgeofFirstMarriage	Coef.	Std. Err.	t	P>|t|	[95% Conf.	Interval]
Post_Reform	.2445827	1.076419	0.23	0.821	-1.913507	2.402673
proportionOfRespondents	.1125005	.0105065	10.71	0.000	.0914361	.1335648
estimator	.0116783	.0153009	0.76	0.449	-.0189983	.0423548
_cons	11.74531	.7848209	14.97	0.000	10.17184	13.31878

Figure 1: Regression of Age of first marriage against Post_Reform, proportionOfRespondents, and the difference-in-difference estimator

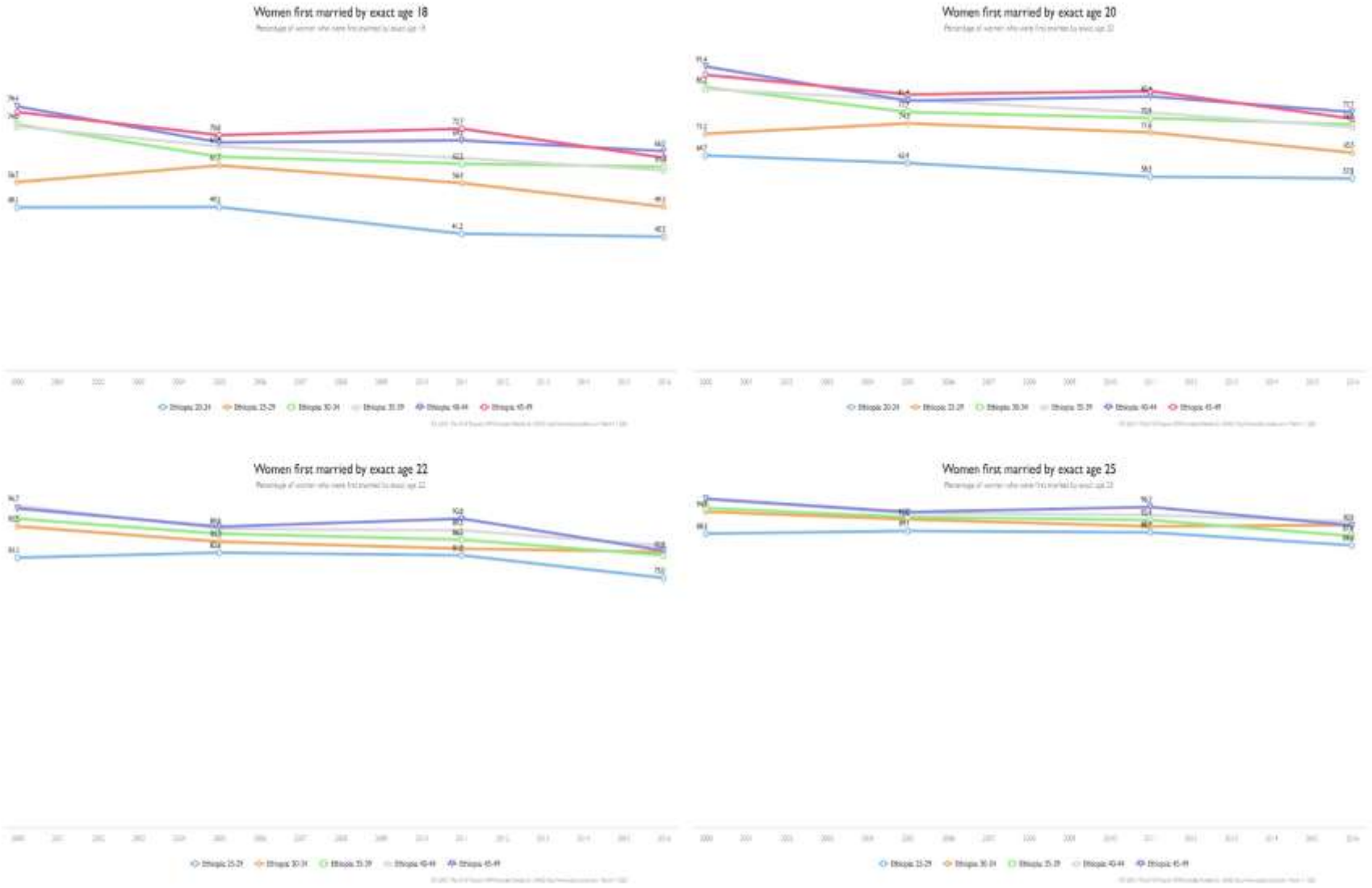

Figure 2: Graphs that review the changes in proportion of women married by ages 18, 20, 22, and 2