SON PREFERENCE: A REASON FOR SLOW FERTILITY TRANSITION IN PAKISTAN?

EXPLORING DURATION MODEL OF BIRTH INTERVALS

INTRODUCTION

Pakistan is the fifth most populous country in the world with 207.8 million inhabitants and a population growth rate of 1.9 percent (Pakistan Bureau of Statistics, 2017). Pakistan is one of the only countries in the South Asian region (excluding Afghanistan) where the fertility rate is still above replacement level (UNFPA, 2017). While fertility has been decreasing in the last decade, the decline has been slow: TFR went down from 4.1 in 2007 to 3.6 in 2018. Additionally, the recent Pakistan Demographic and Health Survey (PDHS) showed that the contraceptive prevalence rate decreased from 35.4 percent in 2013 to 34.2 percent in 2018, making Pakistan the only country in recent history where CPR has gone down (NIPS and ICF, 2019). This has puzzled both researchers and policymakers, especially because the unmet need for contraception is still high, as is the rate of unwanted fertility at one child per woman (NIPS and ICF, 2019). Various reasons have been discussed in the literature for slow fertility transition in Pakistan including poor delivery of reproductive healthcare, religion, social conservatism, and preference for large families (Sather, 2013; Mir and Shaikh, 2013). However, research in Pakistan is limited on the impact of son preference on reproductive behavior.

BACKGROUND

Gender preference in high and low fertility settings

Early literature looking at the impact of gender preference on fertility found no or weak impact. However, most of these studies were conducted in high fertility settings. They highlighted that in high fertility population, gender preference does not impact fertility (Rahman and DaVanzo, 1993). Zaidi and Morgan (2016) maintain that when fertility is high, son preference influences fertility behavior only modestly even in strongly patriarchal societies. In high fertility setting, the lack of association between gender composition of children on fertility was attributed to the fact that with many births, the probability that a woman has at least one son is very high (Zaidi and Morgan, 2016). Son preference increases when families get smaller (Jayachandran, 2015). Guilmoto (2015) argued that decline in fertility has a mechanical consequence, since the risk of not having a child of a given sex increases when the average number of children falls. Only 6 percent of couples do not have a child of a preferred sex when they have four children on average, this risk rises rapidly to 24 percent for those with only two and 34 percent for those with 1.5 children (Guilmoto, 2015). This highlights that once the fertility starts declining, the impact of gender preference needs to be studied with respect to its impact on reproductive decision making.

Effects of gender preference on demographic processes

Son preference exacerbates demographic problems, causing additional births due to son preference and thereby increasing population growth (Zaidi and Morgan, 2016). In the long run, it can also impact marriage rates as an imbalanced sex ratio implies that some men will be unable to find partners in the future (Guilmoto, 2012). Additional births also result in a competition between siblings for the limited family resources (Pande, 2003). Wherein girls are likely to face more discrimination than boys, resulting in lower healthcare and high mortality among girls (Barcellos et al, 2014).

Son preference in Pakistan

Research on son preference is limited in Pakistan. Descriptive results from the PDHS show that girls are at a disadvantage for child mortality, nutrition, and healthcare seeking. Furthermore, in the absence of universal vital registration and census data, it has been challenging to study sex ratios at birth as an indicator of the prevalence of sex selective abortions. In this context, Zaidi and Morgan (2016) use PDHS to study the differential stopping behavior by examining sex ratio at last birth. Their study notes that sex ratio at last birth is high in the range of 120-130, indicating a clear evidence of a behavioral response to sex preference. Their study concludes that Pakistani couples will “continue childbearing to have a son, to have more than one son, and to have at least one daughter” (Zaidi and Morgan, 2016).

In another study Channon (2017) used three rounds of PDHS conducted in 1990–1991, 2006–2007, and 2012–2013 to examine potential indicators and outcomes of son preference. The study found that association of son preference with parity progression and modern contraceptive use has become stronger in Pakistan over time. Channon (2017) further suggested that the prevalence of modern contraceptive use among parous women would have been 19 percent higher in the absence of son preference.

Both Zaidi and Morgan (2016) and Channon (2017) highlighted that son preference impacts fertility and reproductive behavior in Pakistan. Building on their findings, while showing the impact of son preference on fertility and reproductive behavior, my analysis highlights that son preference enforces short birth intervals which has direct consequences for child health and survival. I also show that shorter birth intervals impact child survival which has direct implications for higher fertility rates in Pakistan.

Duration model of birth interval

Rossi and Rouanet (2015) advocate looking at birth intervals to find evidence of gender preferences in a high-fertility context. This model provides a hypothesis that couples will tend to shorten birth spacing to have additional children as long as they have not had enough sons (Rossi and Rouanet, 2015). Using the duration model on North African population, Rossi and Rouanet (2015) infer the existence of son preference when birth spacing is shorter for couples with fewer sons. They deduce that preference for variety prevails when couples having a balanced mix of sons and daughters wait longer than couples having same-sex children (Rossi and Rouanet, 2015). The paper found that son preference influences fertility patterns in North Africa.

The benefit of using duration model of birth intervals is that it accounts for health risks related to birth spacing, and not only to the number of births (Rossi and Rouanet, 2015). For instance, girls are breastfed for a shorter duration because the mother needs to get pregnant again, whereas boys are breastfed for a longer duration because for a lower desire to get pregnant again after the birth of a boy (Jayachandran and Kuziemko, 2011; Barcellos et al., 2014). According to the medical literature on developing countries, short birth intervals are associated with adverse outcomes both for children and mothers (Rossi and Rouanet, 2015). The authors showed that intervals lower than 24 months multiply the risk of infant death by 2.5, and intervals lower than 15 months multiply the risk of maternal death by two (Rossi and Rouanet, 2015). Wide ranging medical literature also shows that short birth intervals are associated with adverse outcomes for both children and mothers (Kozuki and Walker, 2013). Therefore, these short intervals could in turn affect fertility rates by a direct positive relationship between infant mortality and fertility (Mason, 1997; Cleland, 2001; Hossain et al., 2007).

Research question

Does gender composition of the living children influence the duration of the interval before next childbirth?

Hypotheses

Hypothesis-1

Couples with higher proportion of daughters are at a higher risk of having another birth compared to those who have a higher proportion of sons in the family.

Hypothesis-2

Couples with higher proportion of daughters have shorter birth intervals for next birth as compared to those who have a higher proportion of sons.

DATA

I have used the retrospective birth history data from Pakistan Demographic and Health Survey (PDHS) 2017-18. The latest round of PDHS is based on a nationally representative sample of 12,364 married women of reproductive age. Birth history data include information on all live births of sampled women. Using these birth histories, I calculated the birth interval duration from one birth to another. The open birth intervals (duration since last birth at the time of the survey) was considered as censored. The duration for censored cases was calculated by counting the number of months since last birth at the time of the survey. The analysis was limited to birth intervals that were up to 84 months long. This was done to exclude the outliers as well as the cases where women have completed their childbearing many years before the date of interview. The sample was then reduced to 45,547 birth histories. The intervals with a duration of zero (i.e. last child was born in the same month as the month of interview) were excluded from the analysis. This was done for the correct specification of Accelerated Failure Time Model (AFT).

METHODS

I study differential spacing rule using the duration model of birth intervals. The main variable of interest is the duration between births n and n+1, where n $\geq$1. The birth of the next child after the index birth is considered as a "failure" event. Time=0 at the time of index birth. Therefore, for the non-censored cases, the duration to event was at least nine months of pregnancy. For the censored cases, duration could be $\geq$1.

I used Cox proportional hazard regression models to estimate the impact of gender composition of living children on subsequent birth interval. I examine this association in both bivariate and multivariate models. Since Cox models do not predict the survival time, I used the parametric AFT models, to predict the survival time i.e. average duration to the event.

In this analysis, the coefficient of interest is the measure of association between gender composition of living children and the duration to subsequent birth. To measure the construct of gender composition I first calculated a time-varying proportion of sons in the family at the time of index birth. I then classified this continuous measure into the following five categories: all sons, more than half are sons, half are sons, less than half are sons, and all daughters. Panel-I in Table 1 shows the distribution of these five categories in the analytical sample. In panel-II of Table 1, I further classify the gender composition of children in the family into the following three categories: More sons than daughters, equal number, and more daughters than sons. Given the nature of my research question, I will only compare the families that have more sons than daughters to the families that have more daughters than sons. This way, my independent variable is a binary variable and the analytical sample is reduced to 37,153 births.

I have included the following covariates: birth order, death of a child (preceding birth), sex of the preceding birth, mother's age at the time of index birth, mother's education, mother's employment status, wealth status of the household, and area of residence. The selection of covariates was guided by existing literature. Unit of analysis is the child birth, therefore, one woman can contribute multiple births. To account for this, I have clustered the standard errors at mother-level. To deal with the possible problem of ties in rank ordering, I used Breslow estimation methods, which assumes that the cases are tied together due to imprecise measurement.

RESULTS

Descriptive analysis

Average birth interval was 29.1 months long. The analytical sample consists of 37,153 births. At the time of the birth of an index child, 50.9 percent women had more sons than daughters and 49.1 percent had more daughters than sons (Table 1).

Bivariate analysis

The Kaplan-Meier survival curve shows the probability that a study subject survives past a specified time. The KM survival curve presented in Figure 1 shows that around 50 percent of the birth intervals were below 25 months. An important point to note here is that the curve remains straight for the first nine months after which it starts going down, this is because of the inherent nature of the event of interest in this analysis i.e. duration between two births so these nine months represent the pregnancy duration. Figure 2 shows the KM survival curves for both the women with more sons than daughters and vice versa.

Next I examine the bivariate coefficients in the Cox regression model presented in Table 2. The hazard ratio estimate for the women who had more daughters than sons at the time of the index birth were at 6.8 percent higher risk of having another birth compared to those who had more sons than daughters. The hazard ratio estimate of 1.068 is statistically significant at 1%. These bivariate findings are in agreement with hypothesis-1.

Multivariate analysis

Table 2 shows that after controlling for child-level characteristics (birth order, death of preceding child, and sex of preceding child) the hazard ratio for "more daughters" goes up to 1.074 and remains statistically significant (Model-2 in Table 2). Next, I add mother-level covariates: mother's age, mother's education, and her employment status in Model-3. This model shows a further increase in the size of the hazard ratio for mothers with "more daughters" (HR=1.086). Results remain statistically significant. In Model-4, I add the household and community-level covariates: household SES and area of residence (rural or urban). In this last model, the results still show a significant positive association between gender composition of living children and the risk of having another birth. The findings from Model-4 are interpreted as follows: The hazard ratio estimate for the women who had more daughters than sons at the time of the index birth were at 8.6 percent higher risk of having another birth compared to those who had more sons than daughters, after controlling for birth-, mother-, household- and community-level factors. These findings confirm the main hypothesis of this paper that couples are likely to continue childbearing as long as they have not had more sons than daughters.

In Model-4 the hazard ratio for birth order is 0.962 which means that the hazard of another birth decreases as the birth order increases. This is consistent with the existing literature. Death of preceding child increased the risk of having another birth by 40.2 percent. This finding is consistent with the research looking at the impact of child mortality on fertility and reproductive behavior also known as the "replacement effect" of child death. A replacement effect arises when parents experience the death of a child and they consciously or unconsciously change their subsequent reproductive preferences and behavior" (Hossain et al., 2007). Since the analysis did not take into account the timing of death of the child, the given affect size is an under estimate of this association.

The sex of the previous child did not have a significant impact on the risk of having another birth. The size of the coefficient however, was consistent with the existing literature showing that the risk of another birth is shorter after the birth of a son due to various reasons including the fact that boys are breastfed longer than girls (Jayachandran, 2011). Exclusive breastfeeding also affects a woman's fecundability through lactational amenorrhea.

Mother's age was statistically significantly associated with the risk of having another child. The hazard ratio shows that a one year increase in mother's education decreases her risk of having another birth by 3 percent. The women who attended higher than secondary-level education were 19 percent less likely to have another child as compared to the women who had no education. Mother's employment did not have any impact on her risk of having another birth.

The richest households were less likely to have another child as compared to the poorest households, however, this was only significant at 5%. Area of residence was not associated with the risk of having another child.

In a parametric survival analysis AFT model, the dependent variable is logged survival time. Therefore, a positive value of coefficients in Table 3a and Table 3b means that the variable is associated with longer survival, and a negative value means that the variable is associated with shorter survivals. The metric of these coefficients is logged durations in months. The bivariate results in Table 3a show that duration to next childbirth was 35.55 months for the women who had more daughters than sons among their living children compared to 37.23 months among those who had more sons than daughters (Table 3a).

The multivariate AFT model results show that the average duration to next birth was 15.7 months for those who had more daughters compared to the duration of 16.6 months for those who had more sons (Table 3), controlling for child-, mother-, household- and community-level factors. These results are statistically significant at 1%. Furthermore, if a preceding born child died, the average duration to next birth was 13.2 months compared to 16.6 months for those women whose second-last born child was alive.

Heterogeneity analysis

Table 5 to Table 7 show that relationship between proportion of sons and birth interval varies by the background characteristics of the mother.

The hazard ratios presented in Table 4 show that the size of the coefficient was highest for the women from richest families. Where within richest households, those with more daughters than sons were at 17 percent higher risk of childbirth than those with more sons. Similarly, Table 4 shows the evidence for a higher likelihood of “replacement hypothesis” than among any other income strata.

Table 5 shows that women with some education were more likely to manifest their son preference than women with no education. Among women with no education, those who had more daughters were 6 percent more likely to have another child as compared to those with more sons. Similarly, among women with primary and middle level education, those who had more daughters were 20 percent more likely to have another child as compared to those with more sons.

Table 6 further shows that in urban areas the hazard ratio of birth was 1.077 compared to 1.09 for rural areas. This shows that the son preference exists in both rural and urban areas and the magnitude is almost similar.

CONCLUSION

The results of this analysis confirm the preference for sons in Pakistan. The birth intervals are significantly shorter when the proportion of daughters is higher among living children compared to those who have a higher proportion of sons. There is a heterogeneous impact of gender composition of living children on the hazard of next childbirth. This impact varies by household SES and mother's education.

Results of this analysis also confirm the replacement effect of child mortality on fertility. Gender differences in the child replacement have widely been discussed in the literature where a death of a specific sex is associated with increased fertility (Hossain et al, 2007). To test the gendered nature of the replacement hypothesis, the future research will examine if the death of a male child had a significantly greater effect on subsequent fertility than the death of a female child.

Findings of this research has important implications for policy to understand the impact of son preference on fertility and how high levels of child mortality contribute to consistently higher fertility rates.

Tables and Figures:

Table 1: Main variable of interest

<table>
<tr><th colspan="3">Panel-I</th><th colspan="3">Panel-II</th></tr>
<tr><th colspan="3">Gender composition (categorical)</th><th colspan="3">Gender composition (binary)</th></tr>
<tr><th>Classifications</th><th>N</th><th>Percent</th><th>Classifications</th><th>N</th><th>Percent</th></tr>
<tr><td>All sons</td><td>10,217</td><td>22.43</td><td rowspan="2">More sons than daughters</td><td rowspan="2">18,893</td><td rowspan="2">41.48</td></tr>
<tr><td>More than half are sons</td><td>8,676</td><td>19.05</td></tr>
<tr><td>Half are sons</td><td>8,394</td><td>18.43</td><td>Equal number*</td><td>8,394</td><td>18.43</td></tr>
<tr><td>Less than half are sons</td><td>8,979</td><td>19.71</td><td rowspan="2">More daughters</td><td rowspan="2">18,260</td><td rowspan="2">40.09</td></tr>
<tr><td>All daughters</td><td>9,281</td><td>20.38</td></tr>
<tr><td>Total</td><td>45,547</td><td>100.00</td><td>Total</td><td>37,153</td><td>100.00</td></tr>
</table>

*Discarded from analysis, this makes the explanatory variable a binary indicator.

Figure 1a: KM survivor curve estimate for childbirth

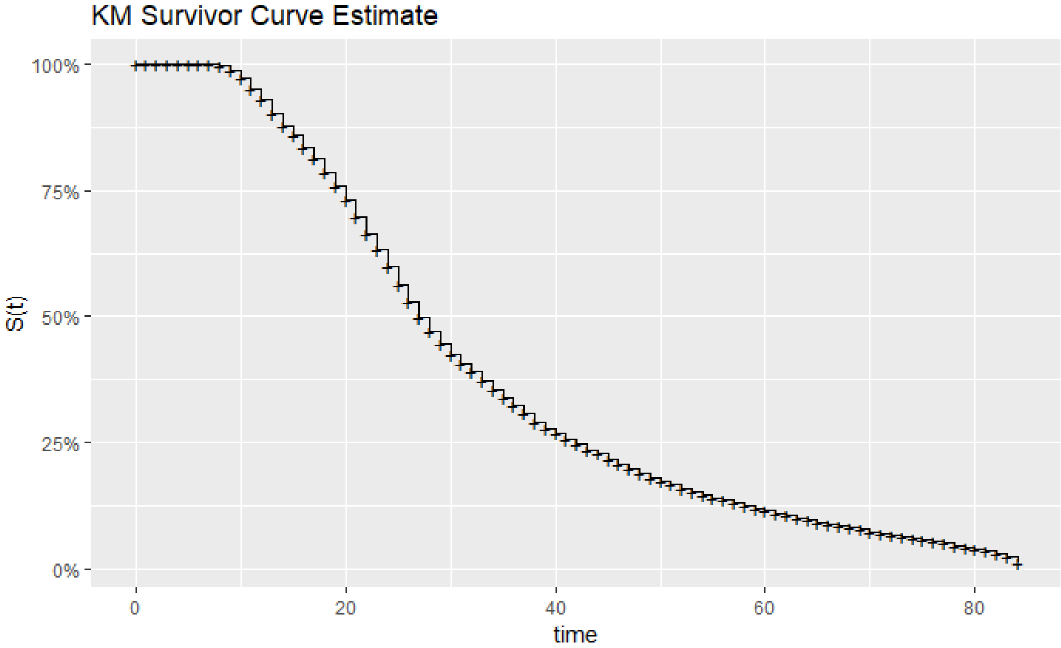

Figure 1b: KM survivor curve estimate for childbirth by gender composition category of the family

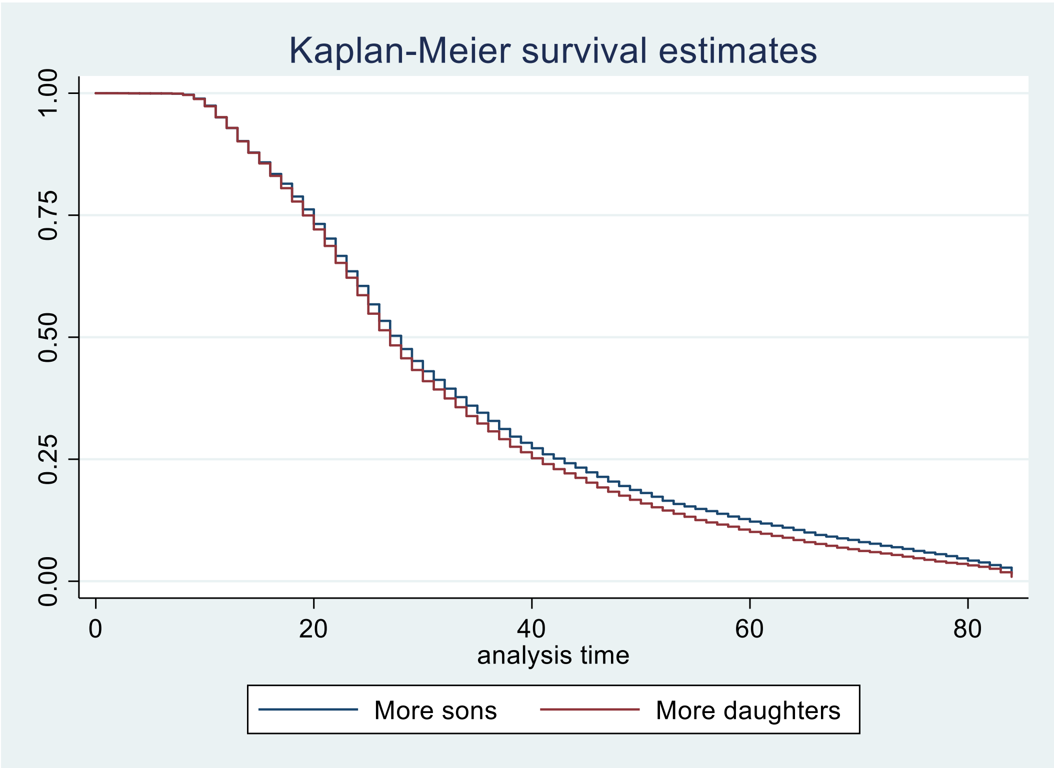

Table-2: Multivariate regression results for the association between gender composition of living children and hazard of next birth

Hazard ratio	**Model-1**	**Model-2**	**Model-3**	**Model-4**
Gender composition of living children				
More sons	1.000	1.000	1.000	1.000
More daughters	1.068***	1.074***	1.086***	1.086***
Birth order		0.916***	0.962***	0.962***
Death of preceding child		1.498***	1.406***	1.402***
Boy (sex of preceding birth)		0.980	0.981	0.981
Mother's age (time-varying)			0.969***	0.969***
Mother's education (no education)			1.000	1.000
Primary			0.913***	0.923***
Middle			0.827***	0.842***
Secondary			0.783***	0.805***
Higher			0.762***	0.794***
Mother's employment status			1.006	1.001
Household SES (Poorest)				1.000
Poorer				0.992
Middle				1.012
Richer				0.981
Richest				0.947**
Area of residence (rural)				0.978
N	37,153	37,153	37,148	37,148

* p<0.10, ** p<0.05, *** p<0.01

Table 3a: Bivariate results from AFT model showing log duration to next birth

Variables	Log T	T	Significance level
More daughters	-0.04618	35.55	***
Intercept	3.61716	37.23	***

Table 3b: Multivariate results from AFT model showing log duration to next birth

Variables	Log T	T	Significance level
More daughters	-0.05821	15.7	***
Birth order	0.02361	17.0	***
Death of preceding child	-0.22963	13.2	***
Sex of preceding birth (boy)	0.01578	16.9	
Mother's age	0.02829	17.1	***
Mother's education (no education)		16.6	
Primary	0.07246	17.9	***
Middle	0.15277	19.4	***
Secondary	0.18972	20.1	***
Higher	0.21330	20.6	***
Mother's employment status	-0.01676	16.4	
Household SES (Poorest)		16.6	
Poorer	0.00464	16.7	
Middle	-0.01389	16.4	
Richer	0.00352	16.7	
Richest	0.00970	16.8	
Area of residence (rural)	0.01314	16.9	
Intercept	2.81214	16.6	***

* p<0.10, ** p<0.05, *** p<0.01

Table-4: Multivariate regression results for the association between gender composition of living children and hazard of next birth by household wealth index

Hazard ratio	Poorest	Poorer	Middle	Richer	Richest
Gender composition of living children					
More sons	1.000	1.000	1.000	1.000	1.000
More daughters	1.057*	1.065*	1.125***	1.093**	1.173***
Birth order	0.976***	0.974***	0.970***	0.934***	0.872***
Death of preceding child	1.378***	1.372***	1.437***	1.405***	1.636***
Sex of preceding birth (boy)	1.026	0.945*	0.976	0.949	1.039
N	8,813	8,824	7,439	6,244	5,828

* p<0.10, ** p<0.05, *** p<0.01
Above models also control for mother's age, mother's education, mother's employment status, and area of residence. Standard errors are clustered at the mother-level.

Table-5: Multivariate regression results for the association between gender composition of living children and hazard of next birth by mother's education

Hazard ratio	No education	Primary & Middle	Secondary & Higher
Gender composition of living children			
More sons	1.000	1.000	1.000
More daughters	1.061***	1.200***	1.145**
Birth order	0.978***	0.927***	0.926***
Death of preceding child	1.341***	1.617***	1.451***
Boy (sex of preceding birth)	0.975	1.035	0.977
N	22,821	4,977	2,710

* p<0.10, ** p<0.05, *** p<0.01
Above models also control for mother's age, mother's employment status, household wealth, and area of residence. Standard errors are clustered at the mother-level.

Table-6: Multivariate regression results for the association between gender composition of living children and hazard of next birth by area of residence

Hazard ratio	**Rural**	**Urban**
Gender composition of living children		
More sons	1.000	1.000
More daughters	1.090***	1.077***
Birth order	0.974***	0.941***
Death of preceding child	1.409***	1.394***
Boy (sex of preceding birth)	1.008	0.945**
N	20,523	16,625

* p<0.10, ** p<0.05, *** p<0.01
Above models also control for mother’s age, mother’s education, mother’s employment status, household wealth, and area of residence. Standard errors are clustered at the mother-level.

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