How Intergenerational Status Transmission Affects Marital Sorting: Evidence from China

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Abstract

This paper studies how intergenerational status transmission influences marital sorting. We find that a policy change in China that gave urban men the same right as urban women to pass on the valuable *hukou* status to their children regardless of their spouse's *hukou* greatly raised their propensity to form inter-*hukou* marriages, compared to unaffected urban women. Additional evidence suggests that a society trapped in a situation with strict marital sorting and little intergenerational mobility for the disadvantaged group can move to a more equitable steady state by reducing the negative impact of social/legal discrimination on intermarriage children. (JEL classifications: J12)

Keywords: Marital sorting, Hukou, Intergenerational status transmission.

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1 Introduction

In his early contributions to the economics of marriage, Becker (1973, 1974) specifies several sources of gains from marriage, among which the importance of raising children is emphasized – "Nothing distinguishes married households more from singles households ... than the presence, even indirectly, of children." If Becker is correct, other things being equal, an individual maximizing his net gains from marriage should select a marriage partner that will produce children of the highest possible quality, instead of simply leaning toward the most beautiful, intelligent, or wealthy partner as implied by the literature on assortative mating (e.g., Fernández, Guner and Knowles 2001, 2005; Wong 2003; Nakosteen, Westerlund and Zimmer 2004).

Even though in theory we can show that concerns regarding offspring can sustain an "aristocratic equilibrium" or "caste equilibrium", in which men and women marry based on inheritable "status" rather than on income/endowment and perpetuate a status-based society (Cole et al. 1992; Banerjee et al. 2009), in practice it is difficult to tell to what extent rigid assortative mating on "status" is due to concern with respect to prospective children rather than to concerns regarding spouses per se because marriage partners' attributes may well correlate with those of prospective children.

This paper estimates the impact of intergenerational status transmission on mate choice by exploiting a quasi-natural experiment in China's *hukou* (household registration) inheritance law.[1] For ordinary Chinese, hereditary *hukou* is the most important determinant of one's socioeconomic status, an effect similar to caste for Indians. Changes to *hukou* inherited from parents are tightly controlled; the official rules are based on attributes valued by the government, such as advanced education or loyal service in the political apparatus, rather

[1]In Becker's theory, children quality is produced by a combination of family goods (like teaching your kids to read, cooking a nice dinner at home, etc) and goods purchased on the market (like clothes, computers, etc). Here we go beyond Becker by looking into social and legal forces which may play an extremely important role in determining prospective children's quality for certain marriages.

than on individual needs to reunite a family or move to take a new job (Wu and Treinman 2004). Thus changes in the inheritance policy may exogenously affect certain children's *hukou* status in a significant way, offering a valuable opportunity for us to study the link between mate choice and status of prospective children.

Unlike the caste system, *hukou* is imposed by the government in order to restrict internal migration, and *hukou* inheritance law is written with a similar purpose. Before August 1998 children in China were granted their mother's *hukou*, a policy meant to deter local urban men from marrying lower-class migrant women. The law was amended in 1998 to give men the same right as women to pass *hukou* to their children. This policy change greatly improved the status of the children of local urban men and migrant women, yet it did not directly affect the status of the children of local urban women and migrant men. Our data also show that the policy change did not have an immediate impact on local urban women's mate choices. We can thus identify the impact of deregulation on inter-*hukou* marriages using the differences-in-differences (DID) model, in which local urban men are considered as the treatment group and local urban women the control group.

Using a 0.095% sample of the 2000 census, we first show that, much like assortative mating on caste in India, assortative mating on *hukou* is pervasive in China, a phenomenon that has received little serious attention in the academic literature. Because the main body of Chinese society is relatively homogeneous and internal migration has been very vibrant since the economic reforms, segregation in the marriage market does not appear to be natural. Our DID estimations show that a light deregulation of children's *hukou* in 1998 raised the interprovincial marriage rate of local urban men by 46% in just two years. Data from the 2005 mini-census further shows that the impact grows stronger as time goes by. In other words, before the policy change local urban men did not want to marry migrant women at least partly because their prospective children would inherit the mother's *hukou*. More generally, our results show that who you choose to marry depends on what kind of children you expect to produce.[2]

A major concern regarding the DID model is that the treatment and control groups differ in unobservable characteristics, and hence would respond to the policy change differently (Heckman 1996; Besley and Case 2000). To address this problem we further checked our results using the changes-in-changes (CIC) approach (or nonlinear DID) proposed by Athey and Imbens (2006), which allows the effects of both time and the treatment to differ systematically across individuals. Our CIC estimations are consistent with those of the DID model.

Understanding the extent to which mate choice is driven by concerns in regard to prospective children is not only interesting in its own right, it also has an important implication for the link between marriage and intergenerational mobility. It is commonly recognized that assortative mating, an important feature of almost all marriage markets,[3] dampens intergenerational mobility and causes persistent social inequality.[4] However, if considerations about the next generation dominate the marriage calculus, the logic can also run the other way, and the result could become more vicious through a multiplier effect. If upper-class people expect social inequality to persist in the next generation, they are unlikely to marry someone from a lower social class, no matter how desirable this person is as a future spouse; this will further deepen social inequality in the long run through a feedback loop until the society reaches a steady state with persistent inequality and rigid marital sorting, i.e., a caste system. A simple model adapted from Kremer (1997) confirms our intuition in a more

[2]According to Kalmijn (1998), marital sorting patterns can also arise due to search frictions. For example, assortative mating on status might simply reflect the fact that people spend much of their time in the company of others with similar social status. Our results are less in favor of this view (also see Hitsch, Hortacsu and Ariely 2010)

[3]Vast evidence has shown assortative mating with respect to the spouse's attributes such as education, religion, income, ethnicity, and social class, etc. For an incomplete list, Scully 1979; Lewis and Oppenheimer 2000; Nakosteen, Westerlund and Zimmer 2004; Restuccia and Urrutia 2004; Schwartz and Mare 2005; Fryer 2007; Banerjee et al. 2009, etc.

[4]Many think marital sorting will lead to inequality. Kremer (1997) shows that this is not necessarily true, though it does affect intergenerational mobility. Fernández, Guner and Knowles (2001, 2005) argue that marital sorting and inequality are endogenously determined through changes in the supply of skilled workers.

formal way.

With widening disparity between people with different *hukou* (e.g., Khan and Riskin 1998) and little inter-*hukou* marriage, Chinese society seems to be trapped in such a "caste equilibrium." Can we move the society to a more equitable equilibrium merely by removing discrimination against intermarriage children? Additional analysis of our data provides a positive answer. When Chinese young people become less worried about their prospective children's *hukou* status, their marriage patterns do appear to deviate away from the caste equilibrium. Local urban men are not only marrying more migrant women, they are marrying women from much lower classes in particular, i.e., those with *hukou* from very poor and/or rural areas. There is also evidence that local urban women are following suit, further contributing to a marriage market with less social segregation.

The remainder of this paper is organized as follows. Section 2 introduces the background. Section 3 describes the census data and uses it to present the migration and interprovincial marriage patterns. Section 4 describes the empirical strategy. Section 5 reports estimation results for interprovincial marriages. Section 6 provides additional evidence that the marriage rate across unequal social classes (rich vs. poor provinces, rural vs. urban areas) was improved by the new *hukou* inheritance law. In section 7 we discuss the implications for intergenerational mobility using a simple model. Section 8 concludes the paper by offering some policy recommendations.

2 Background: *Hukou* and Migration in China

Since the late 1950s during the Great Famine, the national government in China has strictly implemented the *hukou* (household registration) system, mainly to control migration from rural to urban areas (Cheng and Selden 1994). *Hukou* has remained a dominant factor in shaping an individual's socioeconomic status. In the era of plan economy, urban jobs were reserved for local urban citizens and mandatorily allocated through a centralized system. Migration was tightly controlled. To limit migration through marriage, which typically involves migrating women who marry up, the government further required that a child's *hukou* follow that of his or her mother. Changing one's *hukou* was almost impossible for ordinary people. The most important exception was college graduates, who were mandatorily assigned to positions across the country and were accordingly granted the *hukou* of workplaces.[5]

Since the beginning of economic liberalization in the late 1970s, labor markets have gradually replaced mandatory allocation. The development of the private sector has increased the demand for cheap labor. Rural residents crowd into cities to seek opportunities. Despite increased mobility in the labor market, migration with *hukou* did not increase and remained around 0.15% to 0.2% each year following the economic reforms (Lu 2003). Since rural residents are unlikely to obtain an urban *hukou*, they are discriminated against and work under conditions that are unacceptable to urban citizens. They are not entitled to many local privileges, such as inexpensive health care and education. The majority return to their villages after several years of hard work in cities (e.g. Zhao 1999, 2002; de Brauw et al. 2002). The rural-urban disparity is widening in this fast-growing economy (Sicular et al. 2007).

In addition to rural-urban migration, the mobility of college graduates has also increased. In the 1980s when the state-owned sector was still dominant in terms of employment college graduates were mandatorily or semimandatorily allocated to positions. Since state-owned enterprises declined dramatically in the early 1990s, mandatory assignment was not sustainable. Since then allocation has been "marketized." College graduates have to find jobs for themselves upon graduation. Many seek opportunities in cities away from their hometown. Jobs in some large firms or public sectors may be associated with prospects of being granted a local *hukou*, but in general it remains difficult to obtain local *hukou* in big cities for nonlocal

[5] Many of college graduates were allocated back to their home province though. Nevertheless, those from rural areas could get urban *hukou* in this way.

graduates as well.

Despite recent debates on whether the *hukou* system should be abolished, reform of this system has been very slow. A main reform occurred at the end of August 1998 when the national government relaxed the restriction of the *hukou* inheritance law to allow both men and women to pass *hukou* to their children.[6] Changing *hukou* through marriage is still tightly regulated. One cannot easily obtain local *hukou* just because one marries a local resident, and migrant men and women are treated equally in this respect.

3 Marital Sorting on *Hukou* in Urban China: Data and Patterns

In this section we present our sample drawn from the population census and describe the patterns of migration and marital sorting in urban China using the data.

3.1 Data and Measurement

This paper draws on two sources of data. The first is the 0.095% sample drawn from the 2000 Population Census of China; the second is the 2005 mini-census that surveyed about 1.5% of the population. Both data contain the year of first marriages. Our analysis focuses on interprovincial and urban-rural marriages, mainly because across-province and rural-urban inequality is most prominent in China. Also, the measures for these two types of marriages are less noisy. We first focus on interprovincial marriages in this section and the subsequent two sections, and reserve the analysis of rural-urban marriages for section 6.

Given that the census data is cross-sectional, detailed information on individual migration history is unavailable. However, an individual's birthplace is reported in the 2000 census data. Therefore we define an individual as "local" if he or she was born in the same province

[6]Source: State Council Order 1998 No. 24.

as the survey province; otherwise we define the individual as "migrant." We do not consider within-province migration and conservatively base our estimates on cross-province migration.[7] The 2005 mini-census does not contain information on birthplace. Instead, the 2005 survey asks whether one is residing in the same place as five years ago. We construct the proxy for migrant based on this information. The 2005 mini-census can serve to cross-check the estimates generated using the 2000 census data. Moreover, it allows us to explore the effects of the policy change over a relatively long term. For the purpose of our research, we restrict our analysis to local residents who lived in urban areas (including suburban areas) as of the time of the survey, and entered their first marriage during the post-reform period (after 1980).

Our definitions of a "migrant" are likely subject to measurement errors. The definition applied to the 2000 census data may label as migrants some local urban residents who had local *hukou* in the surveyed urban area but were born elsewhere.[8] This definition tends to overestimate the rate of interprovincial marriages. In contrast, the definition applied to the 2005 mini-census may label some migrants as local residents if they have remained in the survey location for more than five years. This definition thus tends to underestimate interprovincial marriages. Comparing the results using the two definitions can give us a better sense of the reasonable range of the rates of interregional marriages and policy impact. Moreover, as this type of measurement errors is unlikely to systematically differ by gender, it should not affect our estimation based on the DID model.

Table 1 presents summary statistics for the 2000 sample. The sample totals 39,512 local residents. Means and standard deviations are reported for individual characteristics, including year of birth, year of first marriage, gender, and ethnicity (minority or not) for the periods 1980-2000, 1996-1997, and 1999-2000 respectively. About 58.5% of the sample

[7] We found that our main conclusions apply to within-province migration as well.

[8] Due to the low mobility in the past 50 years, the probability of one being born out of his *hukou* place is very low. Many researchers have used similar measures to estimate the size of migration population in China (e.g., Johnson 2003).

attained junior high school education or below, 30% held diplomas from high school or vocational school, and 11.5% attained three- or four-year college or beyond. Among local urban residents who married between 1980 and 2000, only 4.4% had spouses born in another province. Among those who entered marriage in 1996 and 1997, 5.8% had spouses born elsewhere; the rate increased to around 6.7% in 1999 and 2000. These numbers can be considered as the upper bound of the estimates for the rate of interprovincial marriages.

Table 2 presents summary statistics for the sample drawn from the 2005 mini-census. To compare with the 2000 data, we divide the sample by years of first marriage. The share of those marrying a migrant is lower than the corresponding number calculated using the 2000 data, about 1% in years 1996-1997 and 1.8% in years 1999-2000. Because we use different definitions for migrants due to data limitation, the estimated interprovincial marriage rate using the 2005 mini-census is about one-fifth to one-fourth of the estimated rate using the 2000 census data. These numbers can be considered as the lower bound of the estimates for the rate of interprovincial marriages.

3.2 Patterns of Migration and Marital Sorting

The estimated interprovincial marriage rate (together with the rural-urban marriage rate discussed in Section 6) seem to be abnormally low. Chinese society is relatively homogeneous, with 91.59% of the entire population in the 2000 census composed of ethnic *han* people. Most people share a very similar language, history, culture, political ideology, and religious belief. A quick look at the data shows that the usual culprits such as geographical or cultural differences cannot easily explain this phenomenon.

First, let us look at the geographical factor. It is well known that migration monotonically increased after economic reforms (Lin et al. 2004). The scale of the massive internal migration within China is unprecedented in human history. Figure 1 shows the share of migrants in the urban population by birth cohort and gender (constructed from the 2000 data).

The out-of-province mobility in both genders exhibits a nearly linear increasing trend. The share of out-of-province male migrants is less than 12% in the 1960 cohort (aged eighteen in 1978, the beginning of economic reforms in China). Each subsequent cohort had an increase of about 0.5 of a percentage point relative to the preceding cohort. In the 1960 cohort only about 8% of urban females were born in another province. In the 1980 cohort (aged twenty in 2000, the year of the survey), about 18% of urban females were born in another province. The share of female urban migrants closely followed the pattern for males.

If marital sorting is due to search costs, it would be natural to expect that the rate of interprovincial marriages also increases over time, following the trend of migration. However, the pattern of interprovincial marriages differs significantly from that of interprovincial migration. Figure 2 shows the percentage of local urban residents marrying migrants across years of first marriages by gender. From 1978 to 1991 the share of local urban females marrying migrants remained almost constantly around 4%, slightly higher than the share for urban males. The late 1990s saw a small increasing trend in this share with small variations from year to year. The percentage of local urban men marrying a migrant followed the same trend before 1999. Yet the trends diverged dramatically following *hukou* deregulation in 1998. In 2000 the percentage of urban males marrying migrant women had a sharp increase of 3 percentage points, or about a 60% increase relative to the year before; while the percentage of urban females marrying migrant men stayed almost constant.

Second, cultural differences do not seem to have played a major role. If they were important they should have affected interethnic marriage much more than interprovincial marriage. Marriages across different ethnic groups in China, which typically involves a husband and wife from more distinct cultural backgrounds, constitute about 3.23% of all marriages in the 2000 census. There is also a clear increasing trend. The cross-marriage rates for China's fifty-five official ethnic minority groups range from 1.05% to 86.96%. The median is about 25%, many times higher than the estimated inter-provincial marriage rate.

In particular, 1.05% of married Uyghurs (the least out-marrying minority), 7.7% of married Tibetans, and 13.3% of married Chinese Muslims (*hui*) married out, often with one from the dominant *han* group (Li, 2004). These are all ethnic minorities who sometimes have difficult or even conflicting relationships with the *han* people. The rate of interprovincial marriage is also low compared to other countries. For example, in the United States, where marrying across racial lines is considered to be a rare event, interracial marriages today account for approximately 1% of white marriages and 5% of black marriages (Fryer 2007; Wong 2003).

Because marital sorting in urban China cannot be easily explained by geographical or cultural forces, we explore the marriage patterns for more clues. We specify the following regression:

$$I(migrant\_spouse_i) = \alpha + \beta X_i + \epsilon_i \tag{1}$$

where $I(migrant\_spouse_i)$ is an indicator for local resident i marrying a migrant, which equals 1 if local resident i's spouse is from another province, and equals 0 otherwise. X_i is a vector of explanatory variables including local resident i's gender, education attainment, ethnicity, year of birth, year of marriage, and local migrant density. For convenience in calculating the marginal effects we use the linear probability model. The results are also robust for non-linear models, such as probit and logit models.

We estimate Equation (1) separately for urban locals who married during the 1980s and the 1990s using the 2000 sample. Table 3 reports the estimates. Columns (1) and (3) report the results for specifications including main covariates for the 1980s and 1990s respectively. We also allow the effects of covariates to vary by gender by including interaction terms between the $Male$ dummy and all other covariates. Estimates for such specifications are reported in columns (2) and (4) for the two periods respectively.

Table 3 uncovers several paradoxical patterns. First, local urban women were more likely to marry migrants in the 1980s while the gender effect disappeared in the 1990s, consistent with graphical evidence in Figure 2. Also, more-educated locals were more likely to marry migrants. The education effects are particularly strong for local women. These results seem to contradict the common belief that less-educated local men are more likely to be on the margin of marrying migrant women in a typical marriage market where "(local) men marry down, and (migrant) women marry up."

This paradox can be better explained in the unique Chinese institutional setting. During the 1980s and the early part of the 1990s, highly educated people in China obtained their jobs through government assignment system. Those assigned to work in a place other than their original *hukou* location were always granted the local *hukou* and had the same economic and political rights as local residents. Since their educational attainment and social status were far above average, it is not surprising that they were attractive in the local marriage market during this period. Moreover, the male-female ratio of college graduates at that time was very biased, on average beyond 1.6 for the 1960-1970 birth cohorts. This ratio is beyond 1.68 for migrants born between 1960 and 1970.[9] It is not surprising that many of the best-educated local women would need to marry out (though not really marrying down). These results suggest that the *hukou* policy is the fundamental force structuring the marriage market.

4 Identification Strategy

To rigorously separate the impact of prospective children's *hukou* from other factors (including other *hukou* factors) that affect inter-*hukou* marriage, we need to go beyond our descriptive analysis in the previous section. The 1998 change in *hukou* inheritance law created a quasi-natural experiment that allows us to credibly identify this impact. Since the policy-making process was not transparent in China, the change was unexpected. The mar-

[9] Calculation is based on our sample of 2000 census.

riage market participants were unable to adjust their behavior in advance. If prospective children's *hukou* status matters in marriage decisions, this policy change would increase the willingness of local urban men to marry migrant women. On the other hand, it would not directly affect local urban women's choice of marriage partners, at least in the short run, since they could always pass their *hukou* to their children regardless of their husband's *hukou* status. We therefore take local urban men as the treatment group and local urban women as the control group and use the differences-in-differences model (DID) as our benchmark.[10]

We first examine the short-run changes in the rate of interprovincial marriages using the 2000 census data. We focus on marriages in 1996-1998 (the pre-treatment period) and those in 1999-2000 (the post-treatment period). Since the *hukou* inheritance law was revised at the end of August 1998, by using the entire year of 1998 as part of the pre-treatment period we implicitly assume that dating behaviors required more than three months to respond to the policy change.[11] By using the following two years as the post-treatment period we implicitly assume that two years are enough for marriage outcomes to reflect changes in dating behavior. These assumptions are realistic in China. According to a large-scale online survey conducted by the most popular Chinese-language website sina.com, 30.3% of the survey respondents married their spouses after knowing them within a year, and 44% did so after one to three years. Moreover, when asked about the optimal length between meeting one another and marriage, 31.4% chose one year, and 45.8% chose one to three years.

Our benchmark model is specified as follows.

$$I(migrant\_spouse_{it}) = \alpha_1 Male_{it} + \alpha_2 post + \alpha_3 Male_{it} \cdot post + Z_{it}\mu^1 + \epsilon_{it} \qquad (2)$$

[10] Policy impact on inter-*hukou* marriage can be measured either at the couple-level or at the individual level. The individual-level estimation under-states the policy influence as it double counts the marriages between local men and local women. That's why researchers usually use the measure based on the couples. Here we use the individual-level measure because we need to use DID model.

[11] All our results are robust to excluding marriages in 1998 from the sample.

where $I(migrant\_spouse_{it})$ is an indicator for local resident i marrying a migrant in year t, $I(migrant\_spouse_{it}) = 1$ if i's spouse is from another province, and $= 0$ otherwise; *post* is an indicator for the change in *hukou* inheritance law, $post = 1$ if i married after 1998, $post = 0$ if i married before 1998. $Male_{it}$ is an indicator for the local resident i being male. Z_{it} is a vector of covariates including individual i's education attainment, ethnicity, and year of birth; the prefecture-level migrant density; as well as provincial dummies.

The key parameter of interest is α_3, the coefficient of the interaction between *post* and $Male$ in equation (2). The term α_3 captures the increase in the rate of interprovincial marriage for local men relative to local women due to the change in *hukou* inheritance law. We expect α_3 to be significantly positive if concern regarding prospective children's *hukou* status enters into the marriage decision.

The DID analysis assumes that the control and treatment group follow the same trend before the policy change (parallel-trend assumption). This can be confirmed in Figure 2. More formally we can look at the coefficients before the interaction of the "Male" variable and the "Year of 1st Marriage" variable in Table 3. Even though the rates of interprovincial marriage remained more or less constant in the 1980s and started to increase in the 1990s, the trends are not significantly different for males and females before 1999. This supports the parallel-trend assumption.

The parallel-trend assumption also means that in the absence of intervention the treatment group would have experienced the same change as the control group. One concern about our identification strategy is the violation of this part of the parallel-trend assumption. When local urban men marry more migrants, local urban women may be forced to marry migrants because there are fewer local men in the marriage market. The data favor our assumptions. The change did not occur immediately for local women (see Figures 2 and 3). In the short run, local urban women make a reasonable comparison group. Moreover, even if local urban women were affected by deregulation in this way, our estimates are biased downward. If our estimate is significantly positive it at least represents the *lower bound* of the policy effect on local urban males' interprovincial marriages.

Another issue with the DID approach is that the identified effect may not be the policy effect if the treatment group differs from the control group in terms of the distribution of outcomes in the absence of the treatment (e.g., Heckman 1996; Besley and Case 2000). For example, in our case the distribution of the unobservable tendency to marry migrants may differ between men and women. In the absence of a gendered policy change, men may be more likely to respond by marrying migrants than women if in general such marriages become more acceptable over time. Although the patterns of interprovincial marriages in the 1980s and early 1990s do not support this view, we need to address it in a more rigorous way.

For this purpose, we adapt the changes-in-changes (CIC) model proposed by Athey and Imbens (2006). The proposed model allows (but doe not assume) the treatment group's distribution of unobservables to be different from that of the control group in arbitrary ways, as long as the distributions do not vary over time within groups. The idea of this estimation approach is as follows. We first use the entire "before" and "after" outcome distributions for local women (the control group) to nonparametrically estimate the change over time that occurred for the local women. Assuming that the distribution of outcomes in the treatment group would have experienced the same change in the absence of the intervention, we estimate the counterfactual distribution for local men (the treatment group) in the post-period. We thus can estimate the effect of the intervention on any feature of the distribution.

Finally, we investigate the relatively long-term impacts using the 2005 mini-census data. To examine the heterogeneous effects across years, we apply the event study analysis. The regression is specified as follows:

$$I(migrant\_spouse_{it}) = \sum_{\tau=-2}^{7} \alpha_\tau D_t^\tau \cdot Male_{it} + \beta Male_{it} + Z_{it}\gamma + \epsilon_{it} \qquad (3)$$

where D_t^{τ} is an indicator for the year, which is τ years from the year of the policy change.[12]

A potential confounding factor for our identification strategy is the increasingly biased sex ratio often attributed to the one-child policy, yet the timing is inconsistent with our results. The one-child policy was introduced in 1979 and became mandatory in 1982. The legal minimum age for marriage is twenty-two for males and twenty for females. The cohort who reached twenty-two in 1999 was born in 1977, well before the full implementation of the one-child policy. The cohort born in 1982 reached the age of twenty-three in 2005. However, very few marry at such an early age in the urban areas. In our sample the average age for a first marriage is 26.5 for males and 24.4 for females, with s.d. approximately four years. Because the sex ratio also shot up around the same time in several Asian societies without a one-child policy, a more convincing explanation of China's high sex ratio seems to be easier access to cheap ultrasonography, which had a rapid increase around 1985 (Meng 2009). According to statistics from the Chinese Children Survey conducted in June 1992, sex ratio at birth increased only slightly, from about 1.10 to 1.125 between 1978 and 1986. The steepest increase came as late as 1987, and within five years the ratio climbed close to 1.25 in 1991. In summary, the rising sex ratio cannot confound our estimations because this phenomenon occurred much later.

5 Impacts on Inter-provincial Marriages

5.1 The Short-run Impacts

Table 4 reports the results of our benchmark DID model (2). Columns (1) and (2) report the results of the basic DID specification without and with control variables respectively. The estimates of policy impact on rates of interprovincial marriages from these two specifications are remarkably similar and are both statistically significant at 5% level. Compared to local

[12] τ being negative means $|\tau|$ years ahead of 1998.

females, the likelihood of local urban males marrying migrants increased by 2.7 percentage points. It is worth noting that the rate of interprovincial marriages remained around 5.8% before the policy change. Our estimates thus show that changing the *hukou* inheritance law raised the likelihood of local men marrying migrants by 46% compared to local women. This is a striking increase within only two years.

The effects can be seen more clearly when we examine local males and local females separately. Columns (3) and (4) of Table 4 report the results for the male and female subsamples respectively. The rate of marrying migrants for local men increased by 3 percentage points after deregulation, while the change for local women is negative 0.3 percentage points and statistically insignificant.

We also use the CIC approach to double-check our estimates. Table 5 lists the results for the effect on local urban males (the treated). We report four estimators: (i) point estimates in the discrete CIC model through covariates, (ii) the discrete CIC model with conditional independence, (iii) the discrete CIC model lower bound, (iv) the discrete CIC model upper bound. For estimators (ii)-(iv), we report the most noisy estimates and do not take into account the effect of covariates. For each estimator we present both the estimates for policy effects and their standard errors based on 1,000 bootstrap replications.

Since there is sufficient variation in covariates, point identification can be obtained for the discrete CIC model. The point estimate for the policy effect on the treated is 2.3 percentage points, slightly greater than the DID estimate. It is also significant at the 5% level. If we do not consider the covariates and assume conditional independence, the estimate for the deregulation effect on the treated is 2.2 percentage points and is significant at the 10% level. The lower bound for the estimate is 2.0 percentage points while the upper bound is 8.43 percentage points. Taken together, even we relax the parallel trend assumption underlying the DID approach, the CIC estimates for the policy effects are of the same magnitude.

Did these local men find their brides in other places and take them to the city, or did they find them among migrants who were already in the city? If the latter is true, the unmet demand for inter-*hukou* marriage appears to be very natural. We do not have enough information to directly investigate this question. However, if the result is driven by the second case, it is expected that the policy effect would be particularly large in regions that already have a strong presence of migrants. Hence we further examine the heterogeneous effects across regions with different migrant densities. We use the following differences-in-differences-in-differences (DDD) model.

$$\begin{aligned} I(migrant\_spouse_{it}) &= \alpha_0 + \alpha_1 Male_{it} + \alpha_2 post + \alpha_3 Male_{it} \times post \\ &+ \alpha_4 Male_{it} \times mig\_density_i + \alpha_5 post \times mig\_density_i \\ &+ \alpha_6 Male_{it} \times post \times mig\_density_i + Z_{it}\beta + \epsilon_{it} \end{aligned} \tag{4}$$

where $mig\_density_i$, the share of migrants in total population in the prefecture where i resides, measures the density of migrants. The coefficient α_6 is expected to be positive if local men respond more to the policy change in areas with more migrants.

Table 6 reports the estimation results of Equation (4). The DDD estimate in column (1) shows that increasing the density of migrants by 1% raised the policy impact by 0.34 percentage points. We further divide the sample into two groups: regions with below-average migrant density and these with above-average density. Columns (2) and (3) list the DID estimates for the two types of regions respectively. Relaxing the children's *hukou* regulation seems to have no significant effect on the rate of interprovincial marriages in regions with a small migrant body. In contrast, it raises the likelihood of local men marrying migrants by 5.2 percentage points in regions with a relatively large migrant body. Compared to the pre-1998 rate of 8.5% for those regions, it represents an increase of 61%. This result indicates that the increase in interprovincial marriages after the policy change mainly occurred between local men and migrant women who were already in the cities. This result also suggests that at least part of the discrepancy between the trend of interprovincial marriage and migration arises from the tight regulation of children's *hukou*.

Do policy effects differ by education level? As mentioned above, college-educated migrants have a better chance of obtaining local *hukou* because some large enterprises offer *hukou* to a limited number of highly skilled employees. With relatively greater options, their marriage prospects are much less constrained by children's *hukou* regulations. We thus expect that college-educated people are less responsive to the change in the *hukou* inheritance law. On the other hand, marriage prospects of migrants without college education are heavily affected by the policy change because they have little chance of obtaining local *hukou* for themselves.

We apply the DDD model similar to Equation (4) (replacing the variable $mig\_density_i$ with low_edu_i, an indicator for educational attainment of less than college) to examine whether the policy effect differs between college graduates and those without college diplomas. From column (1) in Table 7, we can see that the coefficient on the third-degree interaction term ($Male * post * Low\_edu$) is significantly positive, which suggests that the policy change increased the rate of interprovincial marriages at a greater rate among low-educated urban residents than highly educated urban residents. We further decompose the sample into two subsamples – those with a college diploma and those without – and apply the DID analysis to the two subsamples. The results are reported in columns (2) and (3) in Table 7 respectively. Column (2) shows that the policy effect is statistically insignificant for highly educated people, though the sign is positive. Column (3) shows that the policy effect is strong for low-educated people: deregulation increased the rate of interprovincial marriages by 2.8 percentage points, and is significant at the 10% level.

5.2 Temporary or Persistent Impacts?

Given the strong and robust short-run effects, it is worth exploring whether these effects are persistent or temporary. In this subsection we explore the relative long-run effect of the policy, using the 2005 mini-census data.

As mentioned in Section 3, we define locals and migrants using 2005 census in a different way from that used in 2000 census due to the limitation of the data. Yet these two measurements are reasonably close, due to the limited opportunities of migration with *hukou*. Figure 3 shows the rates of interprovincial marriages for each gender by year of marriage using the 2005 census data. The pattern before 2000 is similar to that displayed in Figure 2 where the 2000 data are used. More strikingly, the rates for local men marrying migrants shot up further after 2000, while the rates for local women also jumped, though to a lesser degree.

We estimate the event study analysis model as in Equation (3) using the sample of those whose first marriages fall between 1995 and 2005. We use year 1995 as the reference year. Column (1) in Table 8 reports regression results. Figure 7(a) plots the estimated effects (α_τ) against years. These estimates show that the 1998 policy change tends to increase the likelihood of local men marrying migrants. Compared to estimates for short-run effects (Table 4), the magnitude of estimated effects in 1999 and 2000 is smaller (1.0 and 1.3 percentage points vs. 2.7 percentage points in Table 4). One reason is the aforementioned measurement errors in different specifications tend to bias the estimates in different direction. More importantly, it is quite apparent that the effects tend to increase over time. The estimated policy effects for years 2001 through 2003 range from 2.6 to 4.2 percentage points. After 2003 the policy effect jumped to a level beyond 4.7 percentage points.

In 2005 local urban men had a much higher likelihood of marrying a migrant. Given our previous discussion that the sex ratio is still relatively balanced for young people on the marriage market, this difference is largely due to changes in the composition of the market participants or in their behavior. A closer look at the data shows that the age of first marriage rose much faster after the policy change for local men compared to local women. This is not due to local men delaying marriage. Increasingly, more local men are getting married in each year following the policy change. In 1980 local men's marriages accounted for 50.0% of all marriages involving at least one spouse from the local urban area. Over about the next twenty years the rate rose very slowly to about 50.5% in 1998. After the policy change the rate jumped to about 53.5% by 2005, within just seven years. The extra out-marrying local urban men are thus mainly older bachelors who presumably have searched for years in the local urban marriage market but to no avail. They appear to be at the margin of "marrying down."

6 Is China Shifting Away from the Caste Equilibrium?

Improving the legal status of children from inter-*hukou* marriages not only strengthens local urban men's incentive to marry migrating women, it also encourages two types of marriages– interprovincial and rural-urban marriages — that have the potential to relieve two of the most serious inequalities in China, namely regional and rural-urban disparity. This evidence suggests that China may be able to shift away from the "caste equilibrium" merely by reducing discrimination against children from inter-*hukou* marriages.

6.1 Impact on the Spousal Status Gap

We also expect that this policy change particularly favors migrant brides from poorer provinces. Controlling their virtues, migrant women from poorer provinces were discriminated against more by the old marriage system because the local urban men were concerned that their children would inherit the mother's lower *hukou* status. The policy change regarding children's *hukou* eliminated this concern and made all migrant women equal. This obviously provided a big boost to migrant women from poorer provinces. Other things being equal, we expect to see more marriages between local urban men and such migrant women.

We thus want to investigate the policy impact on the spousal status gap (measured by one's own birth-province GDP per capita minus spouse's birth-province GDP per capita in 2000).[13] There is a huge variation in GDP per capita across province in China. The richest province (with GDP per capita at a level of 34,547 yuan in 2000, with eight yuan roughly equivalent to one U.S. dollar) is about ten times richer than the poorest province (at a level of only 2,662 yuan) in 2000. The median is about 6,500 yuan.

We first examine the short-term effect using the 2000 sample. Figure 4 plots the spousal status gap by marriage-year cohorts, for local urban males and females respectively. We can see that on average both local urban males and females marry someone from a poorer province. The underlying force is easy to understand. Interprovincial migration in China usually occurs when people from poor provinces migrate to more developed provinces (Bao et al. 2007). Thus it is almost certain that local residents who stay in their birth province either marry another local resident, or marry a migrant who most likely comes from a poorer province. The average spousal status gap is very low (about 500-1000 yuan) compared to the provincial difference (on the magnitude of tens of thousands) before 1998. It indicates most marriages occurred between people from the same province (whose status gap is 0 by our measure). Moreover, the trends for local urban males and females remained consistent with each other until 1998. After 1998, however, the spousal status gap for local urban males shot up while that for local urban females stayed stable. This suggests that brides in intermarriages after 1998 had been born in much poorer provinces. The average spousal status gap for local urban men jumped to 2500 yuan in just two years, which is 1.5 times higher than that in 1998. In contrast, there is no obvious evidence that local urban women began to marry migrant men from poorer provinces during the same period.

[13]Source: China Statistical Yearbook 2001. We use GDP per capita in a given year to tease out the effect driven by increasing regional inequality.

To examine the policy effect precisely, we estimate the DID model in Equation (5) using the sample from 1996 to 2000:

$$Spousal\_gap_{it} = \beta_1 Male_{it} + \beta_2 post + \beta_3 Male_{it} \cdot post + Z_{it}\mu^2 + \eta_{it} \quad (5)$$

where $Spousal\_gap_{it}$ measures the gap in birthplace GDP per capita between local resident i and the spouse who married i in year t. The other variables are specified in Equation (2).

Regression results confirm the patterns exhibited in Figure 4. Table 9 reports the results. On average, the spousal gap in birthplace GDP per capita increased by 866 yuan for local urban men relative to that for local urban women. This increase is more than 80% of the pre-reform gap. The result is robust for excluding those who married in 1998. It indicates that when local urban men began to marry migrants more frequently, they were more inclined to marry migrants of a lower status.

The CIC estimates also demonstrate the same pattern. Table 10 reports the results from the continuous CIC models. The CIC estimate for the policy effect on the treated is about 831 yuan. This magnitude is close to that of the DID estimates, and it is also significant at the 5% level. If we treat income as a category variable and estimate the discrete CIC model, the estimates are very similar.

The impact appears to be quite persistent. Figure 5 plots the spousal status gap (defined similarly as in Figure 4) between 1980 and 2005 by marriage-year cohorts using the 2005 census data. The average gap is around 150-250 yuan before 1998. Because our definition of migrants differs across different census analysis, the estimated status gap before 1998 using data from the 2005 mini-census is about one-fourth of the corresponding estimated gap using 2000 census data. This is consistent with the fact that the estimated interprovincial marriage rate in the 2005 mini-census is about one-fourth of the corresponding estimate using 2000 census data. Although the magnitude of the gaps differ due to the use of different measures, the patterns corroborate our above conclusion that bride migration from poor regions dramatically increased immediately after the 1998 policy change. Moreover, the gap for local urban men increased sharply over time, while the gap for local urban women stayed almost constant until 2001 and then experienced an increase as well, suggesting that local urban women are marrying more out-of-province migrants over time.

For more precise measures of policy effects across year, we apply the event study analysis model in Equation (3) to estimate the impact on the spousal gap in *hukou* province GDP per capita. Figure 7(b) plots the estimated effects by year. The estimates are reported in column (2) of Table 8. The spousal status gap increased by about 80% within the two years following the policy change. The effect further jumped by another 40% after 2001.

6.2 Impacts on Rural-Urban Marriages

Studying only the interprovincial marriages in our sample may not capture the major effect on urban-rural marriages if many such marriages occurred between people from the same province. Because such people are closer to each other in terms of geography and culture, the *hukou* barrier to marriages should play a larger role. The 2005 census data allow us to study this complementary research question.

We exclude suburban residents from our sample and focus on those holding urban *hukou* in 2005.[14] We find that about 5.3% of males and 3.9% of females in this sample married rural *hukou* holders of the same province from 1994 to 1997. The pre-change rates are even lower than the pre-change rates of interprovincial marriages. Figure 6 shows the share of local urban *hukou* holders who married rural *hukou* holders across years of first marriage by gender. The shares for both genders exhibit an increasing trend over time. The share for male urban *hukou* holders stays almost constantly around 2% above that for female urban

[14] Although an individual's residence five year ago is known, no information is available on the type of *hukou* (rural or urban) at that time. We thus use the current *hukou* status as a proxy.

hukou holders until 1998, suggesting that a significant urban male fraction chose to marry rural women within the same province despite the *hukou* constraints on their wives and children, and even though over the same period their incentives to marry migrants from another province did not differ significantly from urban women. After 1998 the share for males increased sharply while that for females follows the original trend before 2004. The children's *hukou* policy change appears to have a strong impact on rural-urban marriages.

We can also examine whether rural-urban marriages increase as the result of the policy change, using the event study analysis. The regression specification is the same as in Equation (3) except that the outcome variable is an indicator for local i's spouse holding rural *hukou*. We examine whether the likelihood of marrying a rural resident increased more for urban men than for urban women.

Column (3) in Table 8 reports regression results. Figure 7(c) displays the estimated effects by year of marriage. The figure exhibits similar patterns as Figures 7(a) and 7(b). Compared to the reference year 1995, the likelihood of marrying a rural spouse for urban males relative to urban females remained almost constant in 1996 and 1997. As of 1998, the year of policy change, this likelihood for urban males increased by 1.4 percentage points relative to that for urban females, but it is statistically insignificant. The effects become much stronger and statistically significant from 1999 onward. By 1999 the likelihood of marrying a rural spouse increased by 3.2 percentage points for urban males relative to urban females. Since the pre-reform rate of urban-rural marriages is about 5.3% among local urban men, our estimate shows that the policy change increased the rate for urban men of marrying a rural wife by 60% in one year. The effects for years 2000 through 2002 are of the same magnitude. The estimated effects further jumped around 5.3 percentage points by 2003. This demonstrates that the change in the *hukou* inheritance law has persistent impact on urban-rural marriages. The within-province effects are even stronger than that on interprovincial marriages.

Finally, there is no strong evidence that men from poor and/or rural areas are marginalized by the policy change. According to the long-term pattern shown in our graphs (Figures 3 and 6), local urban women are also marrying more migrant and/or rural men. Even though the increase is not as dramatic as that of local urban men, the increasing trend is clear. Although we do not conduct a thorough welfare analysis, which is difficult in our context, the overall effect appears to be quite positive for most participants in the marriage market, suggesting that desegregation policies created many winners but very few losers. However, the rising sex ratio in the rural areas in particular may work against poor rural men after 2005.

7 A Simple Model of Marital Sorting and Intergenerational Mobility

Many social commentators have raised concerns that marital sorting may reduce intergenerational mobility and exacerbate long-term inequality (Kremer 1997). Across-country evidence shows that countries with low intergenerational mobility tend to have stronger marital sorting (Raaum et al. 2007). Our finding suggests that the causality could also go in the opposite direction. If people are forward looking and care about their offspring, they tend to select partners who are likely to transmit more valuable human/physical capital or better social status to the next generation. The more important the transmission process is, the more likely people would sort by these inheritable characteristics in marriages. Therefore an increase in marital sorting may be the *consequence* rather than one cause of intergenerational persistence in earnings, education, and so on. This channel provides a possible explanation for various patterns of marital sorting across country and over time. For example, Banerjee et al. (2009) find there is a very strong preference to within-caste marriages among middle-class Indians.

We illustrate these points in the simple framework adapted from Kremer (1997). The difference is that we assume that part of marital sorting depends on the expected intergenerational mobility. Without loss of generality, we focus on the transmission of educational attainment. Suppose that the educational attainment of a member of the ith dynasty in generation $t + 1$ is

$$z_{i,t+1} = k_{t+1} + \alpha(\frac{z_{i,t}^F + z_{i,t}^M}{2}) + \epsilon_{i,t+1} \tag{6}$$

where we assume a child's educational attainment is symmetrically affected by his father's and his mother's education, $z_{i,t}^F$, $z_{i,t}^M$ respectively; $\epsilon_{i,t+1}$ is an i.i.d. random shock to educational attainment. We use $corr(z_{i,t+1}, z_{i,t})$ to measure the intergenerational transmission. The term k_{t+1} is an exogenous parameter.

We can use the correlation between spouses' education ρ_t as a measure of marital sorting. Suppose it consists of two parts: one part is dependent on intergenerational transmission; the other part is exogenously given. We thus write it in the following linear form:

$$\rho_t = \rho_m + \gamma E_t(corr(z_{i,t+1}, z_{i,t})) \tag{7}$$

Note that from Equation (6), if spousal correlation ρ_t increases, intergenerational transmission $corr(z_{i,t+1}, z_{i,t})$ will be greater; in turn, the increase in intergenerational transmission further raises marital sorting ρ_t. The term ρ_m is an exogenous parameter. This equation captures the feedback loop between intergenerational mobility and marital sorting. Kremer (1997) takes $\rho_t = \rho_m$ only.

The steady-state correlation between the educational attainment of a parent and his or her child is

$$corr_\infty(z_{i,t+1}, z_{i,t}) = \alpha/2 \cdot (1 + \rho_0)/(1 - \alpha\gamma/2) \tag{8}$$

In the special case where $\gamma = 0$ (no concerns regarding prospective children in choosing mates), the solution in Equation (8) is the same as that in Kremer (1997). An increase of an extra unit in ρ_0 will raise the steady-state intergenerational correlation by $\alpha/2$. However, a positive $\gamma > 0$ will enlarge the effect. Let us define the multiplier as:

$$\phi = 1/(1 - \alpha\gamma/2) \tag{9}$$

The effect of marital sorting on intergenerational correlation can be enlarged by ϕ through the feedback loop between intergenerational persistence and marital sorting. A larger γ always leads to a larger ϕ, suggesting that the multiplier is larger if concerns for prospective children play a more important role in mate selection.

How large can γ be? Our finding suggests that γ is large enough to dramatically increase local urban men's willingness to marry migrants in a short period, but we do not directly quantify γ in the Chinese context. We can borrow an estimate from Ermisch et al. (2006). Using German data, they find that about $40\% - 50\%$ of the covariance between parents' and own permanent family income can be attributed to the person to whom one is married. Although it is not the direct measure of our γ, it suggests a reasonable range. Furthermore, in the developed countries studied by the literature the estimates for ρ_m usually range from $0.5 - 0.65$; the estimates for $corr(z_{i,t+1}, z_{i,t})$ are around $0.4 - 0.6$; the estimates for α largely range between 0.4 and 0.65 (e.g., Solon 1992; Zimmerman 1992; Kremer 1997; Mulligan 1999). Suppose γ is around 0.5 and α is 0.4; the multiplier ϕ will be 1.11 in the overall developed world. If we only focus on racial problems in countries like the United States, the multiplier is likely to be much higher. The multiplier should also be much higher in the poor and segregated societies like India and China.

8 Conclusion

By exploiting a quasi-natural experiment in China, we find that improving the legal status of prospective children from inter-*hukou* marriage greatly improved the willingness of the local urban men to marry disadvantaged migrant women. These findings indicate that to a considerable degree people have already factored their concerns regarding their children's economic prospects into their selection of marriage partners. Our paper helps to explain why inter-*hukou* marriage has been so rare in a relatively homogeneous country with intensive internal migration. In an indirect way, our paper also helps to show the tremendous power of *hukou* in systematically shaping ordinary Chinese people's lives, despite its much-weakened influence after economic reforms.

In other countries where the status of prospective children is largely regulated by social forces, it may be more difficult to credibly estimate the impact of intergenerational status transmission on marital sorting. Even though we suggest readers use caution in generalizing our results, the evidence from the Chinese case is enlightening. If concerns regarding prospective children also play a dominant role in mate choice in India or the United States, we need to approach marital sorting on caste/race differently. Public policies affecting adults may have little impact on marital sorting and long-run inequality. On the other hand, public policies affecting prospective children, in particular those from intermarriage, may be more effective. Because children from intermarriage are implicitly punished by social forces, it makes sense to subsidize the welfare of such children to level the marriage market.

Our simple model on marriage and intergenerational mobility formally shows that concerns in regard to the prospective children always magnify the negative impact of assortative mating on intergenerational mobility in a feedback loop. If we can break this vicious cycle through strengthening certain discriminated groups of prospective children (who may not yet physically exist in our world), society can move to a more equitable steady state with less marital sorting and more intergenerational mobility. The Chinese evidence suggests that this is not entirely unachievable. Increased marriages between men and women from distinct social classes are positive signs of the society slowly recovering from policy-induced social segregation. A small policy change on children's *hukou*, which has received very limited attention thus far, may eventually be remembered as a milestone in breaking the legal barriers to love and intergenerational mobility in China. Further policy changes in this direction are certainly needed to expedite this process.

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Table 1: Summary statistics: sample from the 2000 census

Year of 1st marriage	1980-2000	1996-1997	1999-2000
	Mean(s.d.)	Mean(s.d.)	Mean(s.d.)
	(1)	(2)	(3)
Migrant spouse	.0441	.0581	.0673
	(.205)	(.234)	(.251)
Male	.499	.498	.506
	(.5)	(.5)	(.5)
High school	.3	.313	.383
	(.458)	(.464)	(.486)
College	.115	.19	.223
	(.319)	(.393)	(.416)
Minority	.0364	.0424	.0382
	(.187)	(.201)	(.192)
Birthyear	1963	1971	1973
	(5.79)	(3.96)	(4.22)
N	39512	2737	1782

Notes: Sample includes local urban residents who entered their first marriages between 1980 and 2000 in the 0.095% sample of 2000 census.

Table 2: Summary statistics: Sample from the 2005 mini-census

Year of 1st marriage	1996-1997	1999-2000	2001-2002	2003-2005
	Mean (s.d.)	Mean (s.d.)	Mean (s.d.)	Mean (s.d.)
	(1)	(2)	(3)	(4)
Migrant spouse	.00976	.0179	.0324	.0612
	(.0983)	(.132)	(.177)	(.24)
Male	.501	.504	.51	.512
	(.5)	(.5)	(.5)	(.5)
High school	.259	.274	.28	.295
	(.438)	(.446)	(.449)	(.456)
College	.247	.279	.311	.344
	(.431)	(.449)	(.463)	(.475)
Minority	.0515	.0464	.0514	.0491
	(.221)	(.21)	(.221)	(.216)
Year of birth	1971	1974	1975	1977
	(3.73)	(4.09)	(4.18)	(4.6)
N	12607	10805	8152	10384

Notes: Sample includes local urban residents who entered their first marriages between 1996 and 2005 in the 2005 mini-census.

Table 3: Patterns of Inter-provincial marriages in 1980s and 1990s

	Dependent var.: $I(migrant\ spouse)$			
	Marriages in 1980s		Marriages 1991-1998	
	(1)	(2)	(3)	(4)
Male	-.014	-.714	-.010	4.308
	(.003)***	(1.957)	(.004)***	(4.190)
High School	.016	.026	.008	.017
	(.003)***	(.004)***	(.005)	(.007)***
College	.035	.059	.022	.041
	(.006)***	(.011)***	(.007)***	(.012)***
Minority	-.004	.003	-.034	-.022
	(.009)	(.013)	(.014)**	(.018)
Migrant density	.106	.120	.239	.244
	(.033)***	(.040)***	(.054)***	(.058)***
Year of birth	-.004	-.004	-.006	-.004
	(.0005)***	(.0007)***	(.0007)***	(.001)***
Year of 1st marriage	.004	.004	.006	.006
	(.0007)***	(.001)***	(.001)***	(.001)***
Male*High school		-.020		-.019
		(.005)***		(.009)**
Male*College		-.042		-.035
		(.011)***		(.014)**
Male*Year of 1st marriage		.0008		.00002
		(.001)		(.002)
N	26400	26400	12977	12977
R^2	.026	.028	.04	.041

Notes: Sample includes local urban residents from the 2000 census. Standard errors in parentheses clustered at the prefecture level. All specifications include dummies for province. Specifications in columns (2) & (4) also control for interactions between the $Male$ dummy and minority, migrant density. * significant at the 10% level, ** 5%, *** 1%.

Table 4: DID Results on Inter-provincial Marriages: the Short Run

	Dependent variable: $I(migrant\ spouse)$			
	All		Male	Female
	(1)	(2)	(3)	(4)
Male	-.008	-.013		
	(.007)	(.008)		
post	-.007	.001	.030	-.003
	(.008)	(.009)	(.013)**	(.009)
Male*post	.029	.027		
	(.013)**	(.013)**		
Mig_density		.239	.193	.286
		(.080)***	(.087)**	(.101)***
High School		.009	.015	.010
		(.010)	(.009)*	(.010)
College		.052	-.0008	.055
		(.015)***	(.013)	(.015)***
Minority		-.030	-.047	-.011
		(.021)	(.026)*	(.032)
Birthyear		-.006	-.007	-.003
		(.001)***	(.001)***	(.001)**
Male*High School		.005		
		(.013)		
Male*College		-.052		
		(.019)***		
Provincial dummies	Yes	No	Yes	Yes
N	5795	5795	2899	2896
R^2	0.001	.054	.083	.043

Notes: Sample includes local urban residents who married between 1996 and 2000 from the 2000 census. Standard errors in parentheses clustered at the prefecture level.
* significant at the 10% level, ** 5%, *** 1%.

Table 5: CIC Results on Inter-Provincial Marriages: Short Run

	Effect on the treated group	
	Estimates	(s.d.)
With covariates		
(i) Point estimation	0.0227	(0.0114)**
Without covariates		
(ii) Discrete model: CI	0.0223	(0.0116)*
(iii) Discrete model: lower bound	0.0201	(0.2456)
(iv) Discrete model: upper bound	0.0843	(0.0234)***

Notes: Sample includes local urban residents who married between 1996 and 2000 from the 2000 census. Covariates include migrant density, year of birth, and dummies for province, minority and educational attainment.
* significant at the 10% level, ** 5%, *** 1%.

Table 6: Inter-provincial Marriages: Heteorogenous Impacts across Regions

	Dependent variable: $I(migrant\ spouse)$		
	All Regions	Low Mig_Density	High Mig_Density
	(1)	(2)	(3)
Male	-.014	.001	-.032
	(.010)	(.008)	(.015)**
Post	.011	.006	-.0003
	(.011)	(.011)	(.014)
Male*post	-.003	.0007	.052
	(.015)	(.014)	(.020)***
Mig_density	.201	1.693	.027
	(.111)*	(.358)***	(.112)
Male*Mig_density	.031		
	(.118)		
Post*Mig_density	-.118		
	(.110)		
Male*post*Mig_density	.343		
	(.160)**		
N	5795	3032	2763
R^2	.056	.025	.087

Notes: Sample includes local urban residents who married between 1996 and 2000 in the 2000 census. Standard errors in parentheses clustered at the prefecture level. All specifications include dummies for province and control for minority, year of birth, dummies for educational attainment, and interactions between the $Male$ dummy and dummies for educational attainment.

* significant at the 10% level, ** 5%, *** 1%

Table 7: Inter-provincial Marriages: Heteorogenous Impacts across Education Levels

	Dependent variable: $I(migrant\ spouse)$		
	All Sample	High-education	Low-education
	(1)	(2)	(3)
Male	-.016	-.071	-.011
	(.011)	(.021)***	(.007)
Post	.009	.002	.002
	(.012)	(.034)	(.009)
Male*post	.009	.036	.028
	(.015)	(.030)	(.016)*
Low_edu	-.017		
	(.010)*		
Male*Low_edu	-.009		
	(.011)		
Post*Low_edu	-.016		
	(.016)		
Male*post*Low_edu	.041		
	(.025)*		
N	5795	1171	4624
R^2	.052	.085	.051

Notes: Sample includes local urban residents who married between 1996 and 2000 in the 2000 census. Standard errors in parentheses clustered at the prefecture level. All specifications include dummies for provinces and minority, and control for year of birth and share of migrants. * significant at the 10% level, ** 5%, *** 1%

Table 8: Event Study Analysis - Results based on 2005 mini-census

	Inter-provincial	Spousal status gap	Urban-rural
Dependent var.	$I(migrant\ spouse)$	$Own - spousal\ GDP\ pc$	$I(rural\ spouse)$
α^{τ}	(1)	(2)	(3)
Male* D_t^{-2}	.002	.017	.007
	(.003)	(.030)	(.008)
Male* D_t^{-1}	-.0004	.021	.012
	(.003)	(.041)	(.008)
Male* D_t^{0}	.004	.059	.014
	(.004)	(.061)	(.009)
Male* D_t^{1}	.010	.141	.032
	(.005)**	(.078)*	(.011)***
Male* D_t^{2}	.013	.243	.032
	(.007)*	(.123)**	(.008)***
Male* D_t^{3}	.026	.370	.033
	(.008)***	(.159)**	(.009)***
Male* D_t^{4}	.042	.482	.031
	(.012)***	(.265)*	(.013)**
Male* D_t^{5}	.028	.408	.053
	(.010)***	(.241)*	(.011)***
Male* D_t^{6}	.047	.683	.025
	(.014)***	(.231)***	(.010)***
Male* D_t^{7}	.062	.933	.064
	(.021)***	(.462)**	(.017)***
N	62088	61950	41484
R^2	.049	.059	.05

Notes: Sample includes local urban residents who married between 1994 and 2005 in the 2005 mini-census. Standard errors in parentheses clustered at the prefecture level. All specifications include dummies for province and control for Male, minority, migrant density, year of birth, dummies for educational attainment, and interactions between the $Male$ dummy and dummies for educational attainment. * significant at the 10% level, ** 5%, *** 1%

Table 9: DID Results on Spousal Status Gap: Short run

	Dep. var.: *Own − Spousal birthplace GDP p.c.*(1000*yuan*)	
	Including 1998	Excluding 1998
	(1)	(2)
Male	-.110	-.060
	(.114)	(.132)
Post	-.021	.034
	(.133)	(.120)
Male*post	.866	.815
	(.314)***	(.325)**
N	5769	4497
R^2	.078	.083

Notes: Sample includes local urban residents who married between 1996 and 2000 in the 2000 census. Standard errors in parentheses clustered at the prefecture level. Both specifications include dummies for province and control for minority, migrant density, year of birth, dummies for educational attainment, and interactions between the *Male* dummy and dummies for educational attainment. * significant at the 10% level, ** 5%, *** 1%

Table 10: CIC Results on Spousal Status Gap: Short run

	Effect on the treated group	
	Estimates	(s.d.)
(i) Continuous model	0.8311	(0.3015)**
(ii) Discrete model: CI	0.8284	(0.3015)**
(iii) Discrete model: lower bound	0.8267	(0.3015)**
(iv) Discrete model: upper bound	0.8311	(0.3015)**

Notes: Sample includes local urban residents who married between 1996 and 2000 from the 2000 census. All models include covariates such as migrant density, year of birth, and dummies for province, minority and educational attainment.
* significant at the 10% level, ** 5%, *** 1%

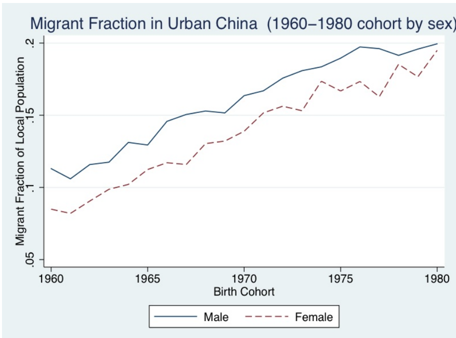


Figure 1: Share of migrants by birth cohort in urban China, 2000

Figure 2: Share of marriages with out-of-province migrants in urban China, 2000

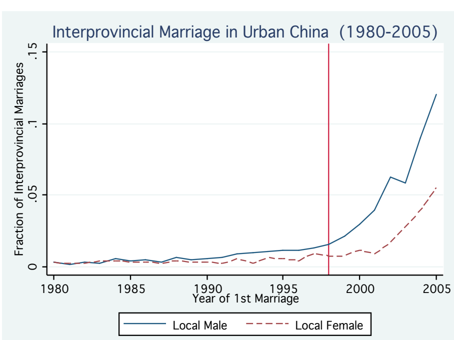


Figure 3: Share of marriages with out-of-province migrants in urban China, 2005

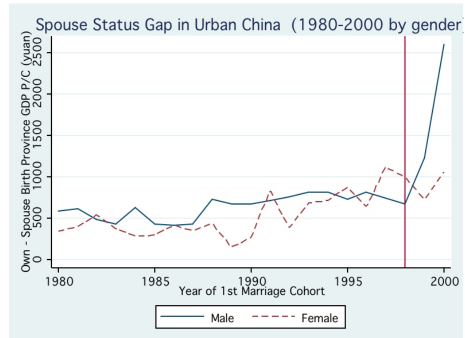


Figure 4: own - spouse birth-province gdp pc gap trend in urban China, 2000

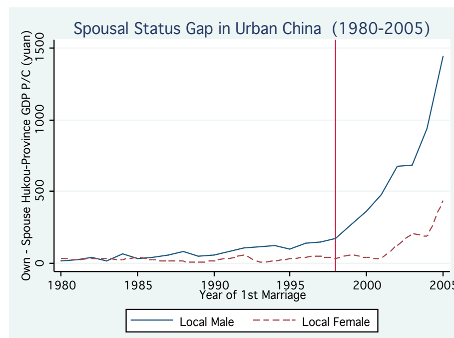


Figure 5: own - spouse birth-province gdp per capita gap trend in urban China, 2005

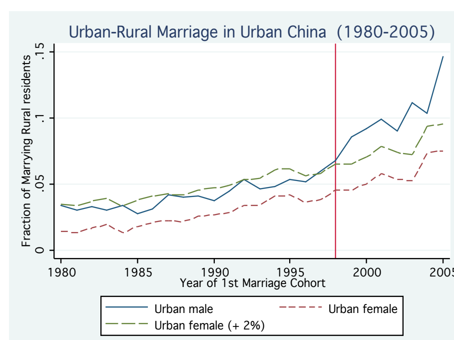
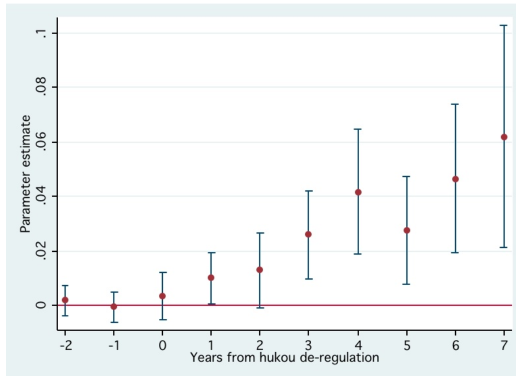
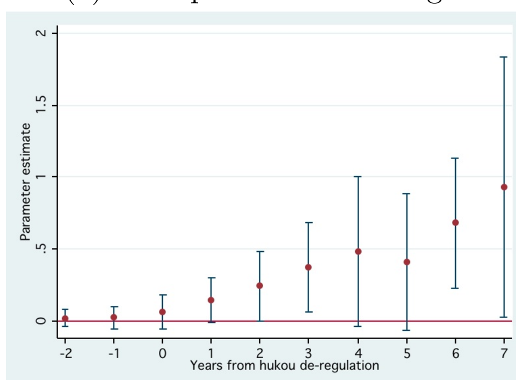


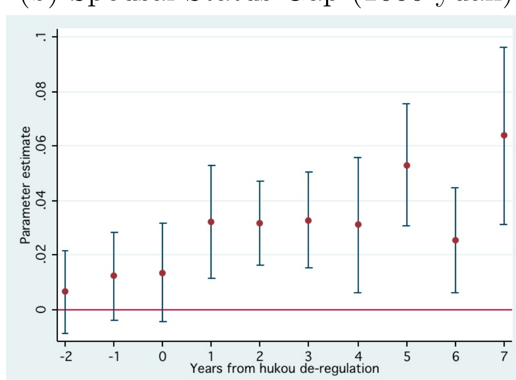
Figure 6: Urban-Rural Marriage trend in urban China, 2005 census



(a) Inter-provincial Marriages



(b) Spousal Status Gap (1000 yuan)



(c) Urban-rural Marriages

Figure 7: Estimates of Policy Impacts from Event Study Models