Markets for Children: International Adoption, ART, and U.S. Foster Care

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Abstract

This paper examines how access to alternative paths to parenthood—international adoption and assisted reproductive technologies (ART)—affects adoption outcomes in the U.S. foster care system. Using a child-level panel from 1998 to 2010, we exploit major shifts in both channels during this period. To address endogeneity, we implement two instrumental variable strategies: a shift-share design leveraging exogenous supply shocks from key international sending countries, and an interactive instrument based on the timing of state-level fertility insurance mandates. Our results show that increased access to international adoption and ART significantly reduces the likelihood that children, especially younger ones, are adopted from foster care. These findings highlight how developments in global adoption policy and reproductive technology can shape domestic adoption outcomes.

Keywords: Foster care; adoption; international adoption; assisted reproductive technology (ART); family economics

JEL Classifications: I38, J13, J18, D12

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1 Introduction

Foster care serves as a critical safety net for children facing abuse, neglect, or other circumstances that prevent their parents from providing adequate care. The system's scale is substantial: in the United States, approximately 5 percent of children enter foster care at some point during childhood, a rate similar to that in other developed countries (Bald et al., 2022). In 2020 alone, the U.S. foster care system held more than 400,000 children, roughly a quarter of whom were awaiting adoption.

While foster care may offer essential short-term protection, its long-term effects on child outcomes remain contested. Early quasi-experimental research documented adverse causal impacts on later-life outcomes such as teen pregnancy and adult crime (Doyle, 2007, 2008). Recent evidence, however, presents a more nuanced picture, with some studies finding positive effects on child safety and educational outcomes, particularly for younger children in more stable placements (Gross and Baron, 2022; Bald et al., 2022). This ambiguity in the literature motivates a deeper examination of the factors that drive positive, permanent outcomes for these children, especially adoption.

For the past three decades, policymakers have relied primarily on financial incentives to boost foster-care adoptions. The Adoption and Safe Families Act of 1997, for example, expanded subsidies and doubled the number of children eligible for support by 2006 (Buckles, 2013; Swann and Sylvester, 2006). While evidence confirms that subsidies can raise adoption rates (Hansen, 2007; Buckles, 2013; Brehm, 2021), the total number of foster-care adoptions remained surprisingly flat for over a decade, increasing only after 2015 (see Figure 1). This persistence suggests that non-financial forces also shape the market for adoption.

This paper investigates two such forces: international adoption and assisted reproductive technologies (ART). During our study period, these alternative paths to parenthood underwent dramatic shifts. International adoptions soared to more than 22,000 in 2004-05 before collapsing to below 2,000 by 2020. Over the same period, ART births quadrupled, rising from roughly 20,000 in 2000 to nearly 80,000 in 2020. As Figure 1 illustrates, these divergent trends motivate our central question: when these other channels become more or less accessible, do prospective parents substitute toward those options and away from adopting children from

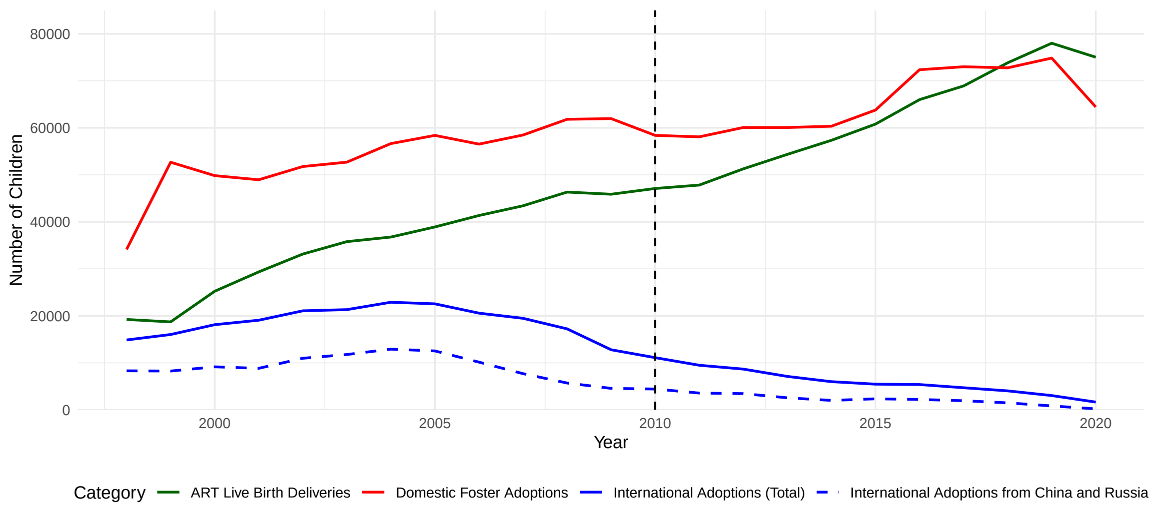


Figure 1: Trends in U.S. Foster Care Adoptions, International Adoptions, and ART Births, 1998-2020

Source: Adoption and Foster Care Analysis and Reporting System (AFCARS), U.S. Department of State, and Centers for Disease Control and Prevention (CDC).

foster care?

To test this hypothesis, we link child-level AFCARS records to state-year counts of international adoptions and ART-birth totals.[1] We use these data in two complementary empirical strategies to identify the causal effects of these alternative family-formation paths.

Our first key finding is that international adoption significantly substitutes for foster care adoption among the younger cohorts. Our analysis focuses on the 1998-2010 panel period, which captures a dramatic rise and fall in international adoptions driven by plausibly exogenous, supply-side restrictions from major sending countries like China and Russia. We exploit this volatility using two methods. Our difference-in-differences (DiD) estimates, which focus on the sharp collapse of these flows, show that the foster adoption rate for toddlers in high-exposure states rose by 64 percent when the international supply contracted. This pattern is confirmed by a Bartik-style shift-share instrumental variable (IV) strategy over the full panel, showing the largest substitution effect for the youngest children (Bartik, 1991; Goldsmith-Pinkham, Sorkin and Swift, 2020; Borusyak, Hull and Jaravel, 2022).

[1]A FOIA request provided disaggregated international-adoption data from source countries to the different states in the U.S. The authors derived the ART live-birth totals for 2007–08 by text-mining CDC clinic-level reports.

In contrast to the volatile trend in international adoptions, ART births grew steadily. To analyze this second channel, we construct an interactive shift-share instrument for ART births that leverages state-level variation in insurance mandates. Our estimates suggest that greater ART access also lowers the adoption probability for foster children, with statistically significant effects for the youngest cohorts. Our point estimates for older foster children are large but imprecisely estimated, warranting caution in their interpretation. However, our finding aligns with the mechanism identified by Zaresani and Schmidt (2024) – that expanded ART access encourages older women to pursue fertility treatments, reducing their likelihood of adoption – a demographic that might otherwise adopt school-aged children.

By jointly analyzing these substitutes with quasi-experimental methods, this paper makes two key contributions to the literature on adoption markets. While prior studies have acknowledged potential substitution (Hansen and Hansen, 2006; Hansen, 2007), few have provided causal estimates that address endogeneity or have jointly considered the role of ART. We provide a broader view of how the demand for foster-care adoption responds to an expanding set of family-formation options.

From a policy perspective, our findings highlight the links in family-formation pathways. Expanded access to international adoption or ART can have unintended social costs by displacing children awaiting adoption in the domestic foster system. These substitutes apply pressure on foster-care adoption rates, underscoring the need to evaluate adoption policy not in isolation, but within the wider context of modern fertility and family-building choices.

The remainder of the paper is organized as follows. Section 2 provides institutional background on U.S. foster care, adoption, and ART. Section 3 describes our data sources. Section 4 details our DiD and IV identification strategies. Section 5 presents the main empirical findings, and Section 6 and Section 7 discuss the implications and conclude.

2 Background

Foster care provides temporary homes for children who have been abused, neglected, or abandoned. While the primary goal is often family reunification, the ultimate objective for children who cannot safely return home is a permanent placement through adoption or legal guardianship (Bald et al., 2022). In 2019, 48 percent of children exiting foster care were reunified with their families, with recent entrants spending a median of 15 months in care. However, lengthy stays are common; 10 percent of the 2015 entry cohort remained in care after four years, often experiencing multiple placements that are associated with poorer behavioral outcomes (Rubin et al., 2004; Bald et al., 2022).

The circumstances leading to foster care entry have shifted over the past two decades, with a substantial increase in cases related to neglect and parental substance use, and a decline in those involving physical or sexual abuse (Bald et al., 2022). A significant driver of this trend has been the drug crisis; studies show that the opioid epidemic, in particular, has increased foster care caseloads and the number of children living with relatives (Buckles, Evans and Lieber, 2023).

There are four primary routes for children to exit the foster system: returning home, aging out of care, entering kinship care or guardianship, and adoption. The legal process for these children is structured around achieving permanency, with family courts overseeing each case. If the court determines that reunification with family is unlikely, the focus shifts to securing a permanent home through adoption, a process that often requires the termination of parental rights (TPR) (Cooper, Doyle and Hojman, 2024; Rashid and Waddell, 2019). This path is formalized through a "case plan," which is determined in dependency court proceedings. A critical milestone is the permanency planning hearing, which must occur within 12 months of placement (Rashid and Waddell, 2019). Between 2000 and 2020, courts assigned a case goal of adoption to nearly 30 percent of foster children and a goal of guardianship to about 2 percent. A child can be assigned this goal before parental rights are finalized; indeed, during our study period, 28 percent of children with a case goal of adoption did not yet have their parental rights terminated (see Table 6). Under 30 percent of children who are assigned an adoption goal were adopted.

In recent years, long-term guardianship with relatives has become an increasingly common permanency outcome, accounting for 17 percent of all exits from foster care in 2019 (Bald et al., 2022). A key distinction from adoption is that for over 95 percent of children exiting to guardianship, parental rights are not terminated. Both adoption and guardianship offer stable, long-term homes, so our analysis considers them permanent exits from foster care.[2]

The broader U.S. adoption landscape is composed of three distinct markets. Public agency adoptions from foster care are the largest component, accounting for 53 percent of all adoptions in 2019 (Gateway, 2022). Intercountry adoptions have seen a dramatic decline, falling from a peak of over 22,000 in 2004 to less than 3,000 in 2019, driven largely by supply restrictions from countries like China and Russia (Shuman and Flango, 2013). The remainder consists of relatively stable numbers of private domestic and independent adoptions (Gateway, 2022).

These different paths to parenthood vary significantly in their costs, demographics, and technologies, as detailed in Table 7 in the Appendix. Adoptions from foster care are the least expensive option and are often supported by post-adoption subsidies (Argys and Duncan, 2013). In contrast, private and international adoptions involve substantial costs, typically ranging from $20,000 to over $40,000, and tend to attract parents with higher incomes and education levels (Moriguchi, 2012).

The costs of ART are also substantial; a single IVF cycle can cost up to $15,000, and total healthcare expenses can escalate to over $400,000 for multiple births (Table 8). High costs have been a major focus of policy, with an increasing number of states mandating insurance coverage of infertility treatments, which in turn increased ART usage (Zaresani and Schmidt, 2024). Moreover, the efficacy of ART improved steadily, making it more viable as an alternative to adoption. For instance, CDC data show that the percentage of ART cycles resulting in a live birth increased from approximately 25 percent in the late 1990s to over 30 percent by 2010. This rising success rate, coupled with lower costs in some states, contributed to the rising births using ART (Gumus and Lee, 2012).

[2] Our main outcome variable, adopted, treats both as adoption. The proportion of long-term guardianship among adoptions rose from 10% in 1998 to almost 20% in 2010.

3 Data

To test our substitution hypothesis, our empirical analysis links three key data components: (i) child-level administrative records from the U.S. foster care system; (ii) state-by-year counts of international adoptions, disaggregated by the child's country of origin; and (iii) annual state-level data on ART births, supplemented with key policy and economic covariates. We construct a child-level state-by-year panel for the period 1998-2010. This window captures the sharp reversal in international adoptions from key sending countries, particularly China and Russia. This volatility, driven by plausibly exogenous supply-side restrictions, provides the variation necessary for our identification strategy. The following subsections detail each data source.

Foster Care

Individual-level data on children in the U.S. foster care system come from the Adoption and Foster Care Analysis and Reporting System (AFCARS). AFCARS provides detailed child attributes, including age, gender, and race, as well as indicators for disability, exposure to abuse or neglect, and reasons for entry into the foster system. We begin our analysis in 1998 due to significant improvements in AFCARS data quality and completeness leading up to that year, a period when the federal government began penalizing states for incomplete reporting (Gumus and Lee, 2012).

International Adoptions

We compile state-year counts of foreign-born adoptees from administrative records of the U.S. Department of Homeland Security, which include the child's country of origin, year of entry, sex, and age group (0, 1- 4, and 5+). A Freedom of Information Act (FOIA) request provided the complete origin-to-destination-state matrix for each year within our sample period.[3] This disaggregated matrix is critical for our empirical strategy, as it allows us to define treatment and control groups for the DiD analysis and to construct the state-specific

[3] This dataset was only available until 2011, which is another reason for our period of analysis.

exposure shares and leave-one-out shifts for our shift-share instrument.

ART Births

Data on assisted reproductive technology (ART) live births are sourced from the CDC's annual reports. As these reports were unavailable for 2007 and 2008, we reconstructed state-level counts for those years by text-mining individual clinic-level data.[4]

Supplementary Data

Our analysis incorporates several state-level controls. Economic and demographic data, including personal income, educational attainment, poverty rates, median household income, and unemployment rates, are drawn from the Bureau of Economic Analysis (BEA), the U.S. Census Bureau, the Small Area Income and Poverty Estimates (SAIPE) program, and the Bureau of Labor Statistics (BLS). In addition, we control for state-level policies that may influence family-formation decisions. We include indicators for triplicate prescription programs for opioids (Alpert, Powell and Pacula, 2018; Buckles, Evans and Lieber, 2023), state mandates for fertility insurance coverage (from the National Conference of State Legislatures), and the presence of state-level adoption tax credits, based on data from Henry and Walker (2023).

Summary Statistics

Table 5 presents summary statistics for the state-year panel. As shown in Table 5, the average annual number of foster adoptions for children under five and international adoptions (which exceed 90% for children under age 5, see Table 7) are of comparable magnitudes, while ART births are already the most numerous during this period.

Table 6 provides summary statistics for the child-level AFCARS data, for children up to age 12 with a case goal of adoption. This group is 52 percent male, 37 percent White, and 26

[4] We thank Can Celebi for compiling the 2007 and 2008 ART live-births data, which we have made publicly available.

percent African American. The data reveal the challenges these children face: neglect is a factor in 64 percent of cases, parental drug abuse in 28 percent, and physical abuse in 17 percent.

4 Empirical Strategy

Guided by a simple choice model (see Appendix for details), we treat international adoption and ART as demand-side substitutes for adoption from foster care. When either channel becomes cheaper, some parents are predicted to switch away from foster children. Because the true costs of those channels are unobserved, we proxy with their equilibrium quantities: more international adoptions or ART births signal lower effective cost. The prediction is therefore straightforward—larger flows of either substitute should reduce foster-care adoptions.

4.1 Model Specification

We estimate how these two substitutes affect the probability that a foster child is adopted. Let $Adopted_{ist} = 1$ if child i in state s exits foster care in year t via adoption or long-term guardianship, and 0 otherwise. A latent "adoptability" index A^*_{ist} drives this outcome:

$$Adopted_{ist} = \begin{cases} 1, & \text{if } A^*_{ist} + u_{ist} \geq 0 \\ 0, & \text{if } A^*_{ist} + u_{ist} < 0 \end{cases}$$

where u_{ist} is a random error and the adoptability index is a linear function of our key variables and controls:

$$A^*_{ist} = \beta_0 + \beta_1 \text{International}_{st} + \beta_2 \text{ART}_{st} + \beta_3 X_{ist} + \beta_4 Z_{st} + \tau_s + \tau_t$$

Here, $\text{International}_{st}$ and ART_{st} are the rates of international adoptions and ART live births per 1,000 women of child-bearing age (15 - 44) in state s and year t. The model includes child characteristics (X_{ist}), state-year covariates (Z_{st}), and state (τ_s) and year (τ_t) fixed effects. Because the majority of adoptions occur among children with a case goal of adoption, our baseline analysis restricts the sample to this group.

4.2 Identification Challenges

A primary challenge in estimating this model is the endogeneity of the key regressors, $\text{International}_{st}$ and ART_{st}. Unobserved factors could bias our estimates of the substitution effects. For instance, a localized economic boom or a shift in social attitudes could simultaneously increase demand for all forms of family-formation (adoptions from foster care or international sources and ART). Conversely, unobserved changes in the "adoptability" of the foster child pool could dampen foster demand while increasing demand for substitutes.

A second challenge is the omission of the unobservable costs of foster adoption. These costs are highly correlated with child attributes and the available supply of foster children (Skidmore, Anderson and Eiswerth, 2016; Gumus and Lee, 2012). We mitigate this by controlling for a rich set of observable child characteristics (including gender, race, disability, and abuse history) and state-level foster population supply.

Finally, our model omits private and independent adoptions. This exclusion is a practical necessity due to the lack of reliable data. As Gateway (2022) notes (p.15), "neither agencies nor attorneys have requirements, incentives, or places to report" such adoptions, a data limitation that is well-documented in the literature (Shuman and Flango, 2013). This omission is also motivated by our theoretical framework (see Appendix), which suggests that the primary substitution effects on foster outcomes will be driven by the international and ART channels. Our identification strategy, detailed below, uses a quasi-natural experiment and instrumental variables to address these concerns.[5]

[5] In the six states where court-record data allow us to estimate private adoption trends, we find that our instrument is orthogonal to these trends—providing suggestive evidence that excluding private adoptions does not bias our main results.

4.3 Identification of Substitution Effects

4.3.1 Difference-in-Differences Design

Our initial approach to identifying the causal effect of international adoptions uses a difference-in-differences (DiD) framework. This strategy leverages a quasi-natural experiment: the sharp and largely unexpected decline in the supply of adoptable children from key sending countries, particularly China and Russia, which began around 2004-2005. Our DiD design leverages the differences in international adoption patterns of high- vs low-exposure states following the supply shock, as visualized in Figure 2. In our primary specification, we compare changes in foster adoption outcomes between the five U.S. states with the highest pre-shock adoption rates from China (the treatment group) and the five states with the lowest rates (the control group), using the shock as the source of identifying variation.

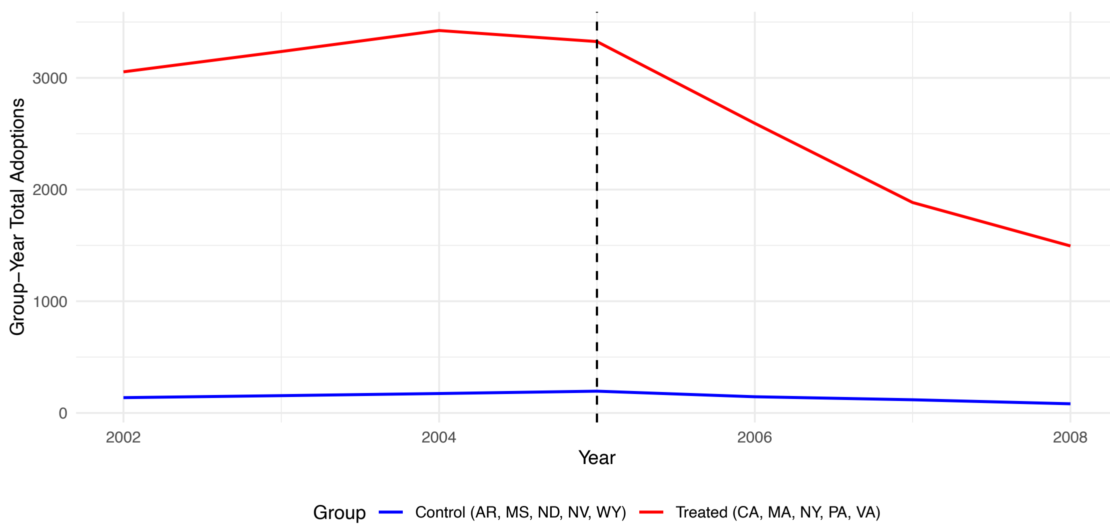


Figure 2: International Adoptions from China and Russia for High-Exposure (Treatment) vs. Low-Exposure (Control) States, Before and After the 2004–2005 Supply Shock.

Notes: The treatment group consists of the five states with the highest pre-shock adoption rates from China and Russia: California, New York, Pennsylvania, Massachusetts, and Virginia. For a China-only analysis, Washington replaces Virginia in the treatment group. The control group consists of the five states with the lowest pre-shock rates. The vertical dashed line indicates the 2004–2005 shock period. A figure showing the nearly identical trends for adoptions from China-only is presented in the Appendix (Figure 5).

Our main DiD estimation examines this effect on the probability of adoption for foster children aged one to six, using a balanced four-year window around the China shock (2001-

2004 versus 2005-2008); robustness checks using three- and five-year windows for China (in 2004) or using the Russia shock (in 2005) yield similar results. Additionally, we conduct placebo DiD estimations to assess pre-trends between the treated and control groups.

4.3.2 Shift-Share Instrumental Variables

To address endogeneity for both substitution channels over the full 1998–2010 panel, we employ shift-share instrumental variables. This approach builds on the literature originally developed to study the effects of immigration on local labor markets (Bartik, 1991) and more recently extended to a range of empirical settings (Goldsmith-Pinkham, Sorkin and Swift, 2020; Borusyak, Hull and Jaravel, 2022). Our application is analogous: we examine the impact of international inflows—adopted children, in this case—on local market outcomes, specifically foster care adoptions, where the inflows themselves may be endogenous.

International Adoptions. A state's volume of international adoptions may be endogenous to its foster care conditions. Our disaggregated data on adoption origin and destination allow us to construct a shift-share instrument composed of two components: historical state-level "shares" from source countries and aggregate, national-level "shifts."

The **shares** (α_{cs}) capture the durable, historical patterns of international adoptions across the U.S., measured as the fraction of all adoptees from a source country c who were placed in state s in a pre-period year (1997). These initial shares we argue are the result of deep-rooted "adoption pipelines" established long before our study period. They are driven by factors such as the historical relationships between the source country and destination state, for example the founding locations of specialized international adoption agencies (e.g., Holt International's long-standing focus on Korean adoptions and its base in Oregon), and the development of local ethnic and religious community networks that lower the costs of adoption from specific regions.

The economic logic of this component is that networks lower the costs of adoptions from specific countries. For example, if a state has a strong history of adoptions from China, local networks will lower the costs of more adoptions from China. The shares measure the differential exposure to common shocks for different states in the U.S. from changes in the availability of adoptive children from China. This framework, where identification is achieved by studying how regions with varying pre-determined exposures react to a common shock, is what Goldsmith-Pinkham, Sorkin and Swift (2020) formalize as an "exposure research design." In our context, the research design asks whether states that were more exposed to a country like China (as measured by their 1997 adoption share) experienced systematically different changes in their foster adoption adoption outcomes following a national-level shock to the supply of children from China.

The **shifts** (Y_{ct}) capture the aggregate, leave-one-out national flow of adopted children from each source country c to the United States in year t.[6] These shifts are plausibly exogenous, driven by supply-side shocks such as the sharp decline in adoptions from Russia following increased regulatory restrictions in 2004, and from China after the imposition of stricter eligibility requirements in 2005.[7]

Instrument Validity. Our instrument, Predicted Adoptions$_{st} = \sum_c(\alpha_{cs} \times Y_{ct})$, interacts the historical shares with the national shifts to predict a state's level of international adoptions. Its validity rests on two standard conditions.

Relevance. The instrument must strongly predict the endogenous variable. States with well-established pipelines from a country like China have specialized agencies and social networks that facilitate these adoptions; when a national-level shift occurs, these high-share states are disproportionately affected. We find a strong correlation between our instrument and actual international adoptions, with first-stage F-statistics that comfortably exceed the conventional threshold of 10.

Exclusion Restriction. The instrument must be uncorrelated with the error term, affecting domestic foster adoptions *only* through its effect on international adoptions. Our

[6] "Leave-one-out" refers to the construction of the shift for each state as the national total minus that state's own contribution. This reduces mechanical correlation with the endogenous variable and strengthens the exogeneity of the instrument.

[7] The U.S. suspended adoptions from Guatemala after it ratified the Hague Convention in 2008, citing systemic non-compliance. While the Convention may not be the sole cause, the complete collapse in adoptions suggests a strong demand-side component. We exclude Guatemala from the DiD analysis for this reason, but include its flows in the construction of the IV, where the identifying assumptions differ.

argument for validity relies on the conditional exogeneity of the initial shares (Goldsmith-Pinkham, Sorkin and Swift, 2020). We argue that the 1997 distribution of adoption pipelines is a historical artifact unlikely to be correlated with unobserved, time-varying factors that influenced foster care adoptions in the 2000s, particularly after controlling for state fixed effects (which absorb all time-invariant state characteristics) and our battery of state-level policy variation and observed confounders such as unemployment and poverty.

This identification is substantially strengthened by the exogeneity of the shifts (Borusyak, Hull and Jaravel, 2022; Jaeger, Ruist and Stuhler, 2018). The major shifts in adoption flows were driven by policy decisions made within sending countries for their own domestic reasons, which were exogenous to foster care conditions in any particular U.S. state. For the exclusion restriction to be violated, one would have to believe, for instance, that the Chinese government's decision to restrict adoptions had a direct effect on Ohio's foster adoptions through a channel other than reducing the number of Chinese children available to Ohioan families. We argue this is implausible, as the shocks were specific to the market for international adoption.

ART Births. We construct an analogous shift-share instrument to identify the causal effect of Assisted Reproductive Technology (ART) live births. The instrument interacts a state's historical **share** of ART activity (its proportion of total U.S. ART live births in 1997) with the national **shift** (the aggregate, leave-one-out trend in ART usage). However, the national trend in ART usage is a smooth, secular increase. As Jaeger, Ruist and Stuhler (2018) highlight, the high serial correlation of such a shift component can make an instrument difficult to distinguish from smoothly trending confounders.

To address this, we introduce policy-driven variation by creating an *interactive* shift-share instrument. We generate sharp, identifiable structural breaks by interacting our baseline leave-one-out Bartik instrument with a dummy variable indicating whether a state has a fertility insurance mandate in effect. This interaction creates distinct changes in the instrument's value for the three states that enacted such mandates within our 1998–2010 panel: **Connecticut** (2005), **New Jersey** (2001), and **Texas** (2005).[8] The plausibly exogenous timing of these

[8]To capture broad support for fertility policies, we code both mandates to offer and mandates to cover

policy changes provides a source of variation that is less likely to be correlated with other confounding trends.

This approach strengthens our identification strategy in two key ways. First, it sharpens the **exposure design** framework of Goldsmith-Pinkham, Sorkin and Swift (2020), as states with different historical ART capacities are now differentially exposed to the interaction of the national trend and a state-specific policy change. Second, by generating distinct structural breaks, our modified instrument is no longer highly serially correlated, enabling the application of the estimation strategy proposed by Jaeger, Ruist and Stuhler (2018).

The validity of this interactive instrument rests on the standard IV assumptions of relevance and exclusion. Relevance is satisfied because a state's initial ART capacity is amplified by insurance mandates, which are known to increase ART utilization. The exclusion restriction requires that the instrument is uncorrelated with unobserved determinants of foster adoptions; this depends on the exogeneity of its three components: the historically-determined 1997 shares, the leave-one-out national shift, and the timing of the state-level insurance mandates. We argue that the timing of the states' decisions to mandate infertility insurance is not directly related to their foster-care adoption rates. For the restriction to be violated, one would have to believe that the legislative decision to enact a fertility mandate in a given year is directly correlated with unobserved factors that are also affecting foster adoptions. We find this unlikely, particularly after conditioning on state and year fixed effects and a comprehensive set of covariates.

4.4 Dynamic Shift-Share Specification and Estimation

Following Jaeger, Ruist and Stuhler (2018), who show that conventional shift-share estimates can be biased by dynamic adjustments to past shocks, we employ a dynamic specification for both substitutes. This model includes a lagged value of the endogenous variable, as shown in the general form:

as a policy being in place. For instance, Texas law mandates that insurers offer IVF plans, but coverage is not required. This differs from studies like Zaresani and Schmidt (2024), which examine only the stricter "mandate to cover" laws.

$$\begin{aligned}\text{Adopted}_{ist} = \beta_0 + \theta_1 \text{Substitute}_{st} + \theta_2 \text{Substitute}_{st-1} \\ + \beta_4 X_{ist} + \beta_5 Z_{st} + \tau_s + \tau_t + u_{ist}\end{aligned} \tag{1}$$

where Substitute_{st} represents either the international adoption rate or the ART birth rate, and Substitute_{st-1} is its one-year lag.

The inclusion of lagged endogenous regressors can introduce its own econometric challenges (Imai and Kim, 2019; Dayal and Murugesan, 2023). To address this, we instrument for both the current and lagged endogenous regressors using current and lagged versions of our respective shift-share instruments. In this framework, the coefficient on the contemporaneous variable (θ_1), our variable of interest, captures the short-run substitution effect and is expected to be negative. The coefficient on the lagged variable (θ_2), as emphasized by Jaeger, Ruist and Stuhler (2018), captures the longer-run dynamic response to past supply shocks and improves the consistency of the estimate for θ_1.

5 Results

5.1 Difference-in-Differences Estimates

Our difference-in-differences (DiD) design exploits the China shock to examine changes in foster-care adoption rates, cohort by cohort, comparing the five most China-reliant states (treatment group) with the five least-exposed states (control group). For each age cohort $a = 1, \ldots, 6$, we estimate the following specification:

$$Adopted^a_{ist} = \beta_a \left(Treated_s \times Post_t \right) + \theta_a \, ART_{st} + \boldsymbol{\delta}'_a \mathbf{X}_{ist} + \boldsymbol{\lambda}'_a \mathbf{Z}_{st} + \tau_s + \tau_t + u_{ist}, \tag{2}$$

where $Adopted^a_{ist} = 1$ if child i in state s and year t exits foster care through adoption or guardianship. The treatment indicator $Treated_s = 1$ for states with the highest shares of

Chinese adoptions prior to the 2005 policy change (2001–04), and $Post_t = 1$ for the four post-shock years (2005–08). The coefficient of interest, β_a, captures the effect of the China supply shock on foster adoptions for cohort a, net of controls.

The specification includes two sets of controls: $\mathbf{X}_{ist}$, a vector of individual-level child characteristics (including indicators for male, presence of disability); and $\mathbf{Z}_{st}$, a vector of state-year covariates (e.g., unemployment rate, poverty rate, child population, and foster-care caseloads). State and year fixed effects (τ_s and τ_t) control for time-invariant state characteristics and national shocks.

Table 1: China: Difference-in-Differences Estimates (4-Year Pre/Post) with Effect Sizes

	Outcome: Foster Adoption Rate					
	Age 1	Age 2	Age 3	Age 4	Age 5	Age 6
Estimates						
Coefficient (Treated×Post)	0.123^{**}	0.151^{***}	0.128^{***}	0.091^{**}	0.102^{**}	0.082^{**}
	(0.047)	(0.053)	(0.046)	(0.038)	(0.040)	(0.036)
ART live birth rate	-0.274^{***}	-0.311^{***}	-0.254^{***}	-0.233^{***}	-0.212^{***}	-0.225^{***}
	(0.084)	(0.099)	(0.080)	(0.074)	(0.077)	(0.057)
Observations	32,249	40,509	37,982	34,086	30,707	27,785
Magnitudes (based on pre-period means in treated states)						
Baseline adoption rate[a]	0.1907	0.3116	0.3387	0.3325	0.3270	0.3137
Relative change (%)[b]	64	48	38	27	31	26

[a] Mean share adopted in treated states *before* the China supply shock (4-year window), by age cohort.

[b] Relative change is calculated as: (Treated×Post coefficient / baseline) × 100.

Notes: Standard errors in parentheses are cluster-robust at the *state_year_id* level. $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

The results in Table 1 show that the negative supply shock from China led to substantial increases in the probability of foster adoption, with the largest gains concentrated among the youngest children. For toddlers (Age 1), the estimated 12.3 percentage-point rise represents a 64 percent increase over the pre-treatment adoption rate. The effect diminishes with age; by Age 6, the 8.2 percentage-point increase corresponds to a 26% rise relative to baseline. This age pattern is consistent with the demographic profile of international adoptions, which were heavily concentrated among children under five. Following the sharp decline in young international adoptees after 2005, we observe a corresponding surge in domestic foster adoptions in the same age groups, particularly in the most-exposed states. These findings are robust to alternative three- and five-year windows and to similar analyses using the Russia supply shock (see Table 10 in the Appendix). To examine pre-trends, we implement a placebo China shock in 2001 with the same treated and control groups. We find no statistical differences in these DiD estimates as seen in Table 9 in the Appendix – indicating absence of pre-trends in foster adoption rates.

Our main DiD estimates Table 1 also reveal a consistent, negative correlation between foster adoption and ART live birth rates. While this reflects a strong association, endogeneity may bias the coefficient. We therefore turn to our instrumental variables strategy using the full 1998–2010 panel.

Panel-Data IV Results

Table 2 presents OLS and IV estimates for the age-1 cohort, which we use to motivate our empirical approach for other cohorts. The OLS estimate in column (1) shows a statistically significant negative association between both international adoptions and ART births, and foster care adoptions, suggesting a substitution effect. However, the IV estimate in column (2), which corrects for endogeneity using a shift-share instrument, is more than ten times larger in magnitude (–2.369 vs. –0.205). This implies that the OLS coefficient is biased toward zero, likely due to positive omitted-variable bias—such as unobserved state-level trends that increase both international and foster adoptions.

The IV estimate for ART live births in column (3) is also substantially larger in absolute terms than the corresponding OLS coefficient (–8.756 vs. –0.108), reinforcing the concern that OLS underestimates the substitution effects. In column (4), we estimate a joint IV model instrumenting both adoption measures. While the coefficient estimates remain economically large, we note a potential loss in statistical precision, as reflected in the increased standard errors and reduced first-stage F-statistics, particularly for ART.

Given these identification concerns, we treat the single-instrument IV specifications in columns (2) and (3) as our preferred models. The corresponding first-stage F-statistics—well above conventional thresholds—indicate strong instrument relevance for both international adoption and ART exposure.

Table 2: OLS and IV Estimates for Age-1 Cohort, 1998–2010

	(1) OLS	(2) IV	(3) IV	(4) IV
Panel A: `intadopt_sy_rate`				
Coefficient	-0.205^{***}	-2.369^{***}		-6.648^{***}
(SE: robust)	(0.055)	(0.654)		(1.475)
(SE: FIPS×year)	(0.124)	(1.734)		(7.967)
(SE: state×year)	(0.209)	(2.329)		(12.20)
Panel B: `art_livebirth_rate`				
Coefficient	-0.108^{***}		-8.756^{***}	-3.541^{**}
(SE: robust)	(0.021)		(2.309)	(1.130)
(SE: FIPS×year)	(0.035)		(18.61)	(5.853)
(SE: state×year)	(0.050)		(30.20)	(9.957)
Observations	140,856	134,194	133,076	133,076
First stage identification statistics (F statistics - state×year clustering)				
International adopt (shift-share)		97.98		68.56
ART (shift-share X fertility mandate)			26.79	13.74

Notes: Each column reports estimates from state-year panel regressions of individual foster-child adoption outcomes (age 1 cohort) on international-adoption rates and ART live-birth rates. Standard errors are shown three ways: heteroskedastic-robust; clustered on FIPS×year; and clustered on state×year. IV specifications use shift–share instruments; column (2) instruments `intadopt_sy_rate`, column (3) instruments `art_livebirth_rate`, and column (4) instruments both. All models include the comprehensive set of state-year covariates described in the text and state and year fixed effects. $^{***}p < 0.001$, $^{**}p < 0.01$, $^{*}p < 0.05$.

Table 3 reports these preferred IV estimates for all cohorts from age 1 to 10. The results confirm two key findings. First, instrumenting international adoptions (Panel A), we find large and statistically significant negative effects for ages 1– 4, a weaker but still significant effect at age 6, and economically small, insignificant estimates otherwise. While this age-gradient is qualitatively consistent with the DiD results, the IV estimate represents a Local Average Treatment Effect (LATE) for the full 1998–2010 panel. Specifically, it captures the average substitution effect for states that are more responsive to the exogenous international supply shocks leveraged by the instrument. Because the shift-share design places more weight on states with high historical exposure to the source countries, this LATE should be interpreted as the effect for this responsive group of states.

Table 3: IV Estimates by Age Cohort, 1998–2010

	Age Cohort (years)									
	1	2	3	4	5	6	7	8	9	10
Panel A: International–adoption rate										
Coefficient	-2.369^{***}	-4.119^{***}	-2.459^{***}	-2.566^{***}	-0.586	-2.454^{**}	-1.142	-1.127	-0.429	-0.566
s.e. (robust)	(0.654)	(0.669)	(0.676)	(0.724)	(0.691)	(0.808)	(0.759)	(0.754)	(0.777)	(0.695)
s.e. (FIPS×yr)	(1.734)	(1.893)	(1.588)	(1.513)	(1.277)	(1.652)	(1.387)	(1.350)	(1.254)	(1.183)
s.e. (state×yr)	(2.329)	(2.715)	(2.244)	(2.148)	(1.725)	(2.243)	(1.826)	(1.715)	(1.600)	(1.431)
Observations	134,194	166,114	152,997	135,170	121,472	112,116	105,683	101,048	97,307	92,359
Panel B: ART live-birth rate										
Coefficient	-8.756^{***}	-9.288^{**}	-12.51	-11.68^{*}	-7.454^{*}	-13.06^{*}	-16.21^{*}	-9.299^{**}	-15.07^{*}	-23.23
s.e. (robust)	(2.309)	(3.042)	(7.299)	(5.294)	(3.045)	(6.108)	(6.683)	(3.179)	(6.413)	(14.63)
s.e. (FIPS×yr)	(18.61)	(25.50)	(66.53)	(45.59)	(21.03)	(51.49)	(54.71)	(22.54)	(50.86)	(119.6)
s.e. (state×yr)	(30.20)	(41.52)	(105.9)	(71.95)	(33.05)	(79.85)	(86.80)	(35.21)	(80.65)	(187.3)
Observations	133,076	164,631	151,630	133,983	120,405	111,157	104,770	100,114	96,438	92,359

Notes: Each column reports a separate IV regression for the indicated age cohort. Panel A instruments the international-adoption rate; Panel B instruments the ART live-birth rate. All models include the full set of child controls, state–year covariates, and state and year fixed effects. The two added rows report the implied change in foster-adoption rates for a +1 s.d. change in each regressor (international-adoption s.d. = 0.07; ART s.d. = 0.17), in percentage points with percent of the baseline in brackets. Standard errors above are reported three ways (robust, FIPS×year, state×year). $^{***}p < 0.001$, $^{**}p < 0.01$, $^{*}p < 0.05$.

Panel B of Table 3 suggests that ART births also crowd out foster adoptions, with negative point estimates across all age cohorts. The estimated effects on older children may be consistent with evidence that fertility insurance mandates encouraged older prospective parents—a demographic that might otherwise adopt school-aged children—to pursue fertility treatments (Zaresani and Schmidt, 2024). While the point estimates for ART are consistently negative and often large in magnitude, a key pattern for cohorts beyond age two is that the standard errors also become large relative to the coefficients. This renders the estimates for older children statistically insignificant in many cases and, while the large point estimates are suggestive, their precise magnitudes should be interpreted with considerable caution. Nevertheless, the pattern of statistically significant negative effects for the youngest cohorts is robust to estimation with an IV Probit model, as shown in the Appendix.

5.2 Magnitudes of Substitution Effects

To gauge the real-world magnitude of these IV estimates, Table 4 provides back-of-the-envelope calculations. The results for international adoptions are substantial: a modest increase of 0.01 in the international adoption rate (per 1,000 women of child-bearing age) is associated with approximately 3,500 fewer adoptions for toddlers (Age 1) over the entire 1998–2010 period. Summing the effects for children up to age four, this increase in the international adoption rate implies a displacement of nearly 20,000 foster adoptions.

The point estimates for ART imply even larger substitution effects for the youngest cohort; a similar 0.01 increase (about a 6 percent shock at the mean) in the ART birth rate is associated with nearly 13,000 fewer adoptions between 1998–2010. However, these magnitudes should be interpreted with caution for two reasons. First, as noted, the IV estimates for ART are large but imprecisely estimated for older children. Second, the results are sensitive to how the instrument is specified. For example, broadening the definition of infertility insurance to include states with fertility preservation mandates yields even larger coefficients.

Table 4: Back-of-the-Envelope Effects of Rates on Foster Adoptions (Ages 1–10), 1998–2010

	Totals (1998–2010)		International adoptions		ART live births	
Age	Foster children	Foster adoptions	ME (IV)	Δ Adopt (+0.01)	ME (IV)	Δ Adopt (+0.01)
1	148,032	33,091	−2.369	−3,507	−8.756	−12,962
2	187,188	65,910	−4.119	−7,710	−9.288	−17,386
3	176,945	65,176	−2.459	−4,351	−12.510	−22,136
4	160,034	56,798	−2.566	−4,106	−11.680	−18,692
5	146,390	49,282	−0.586	−858	−7.454	−10,912
6	136,788	43,860	−2.454	−3,357	−13.060	−17,865
7	130,169	39,823	−1.142	−1,487	−16.210	−21,100
8	125,122	36,377	−1.127	−1,410	−9.299	−11,635
9	120,350	33,225	−0.429	−516	−15.070	−18,137
10	114,530	30,308	−0.566	−648	−23.230	−26,605
Totals $\leq$ 4	672,199	220,975		−19,674		−71,176

Notes: "Foster children" and "Foster adoptions" are totals over 1998–2010 within those assigned the case goal of adoption. Counterfactual changes use ΔAdopt = (total children) × ME × 0.01; negative values indicate fewer foster adoptions when the rate increases.

6 Discussion

The causal impact of foster care placement on children is a subject of ongoing debate. While early quasi-experimental research documented adverse long-term outcomes (Doyle, 2007, 2008), new evidence finds positive effects on safety and education, suggesting that outcomes are highly heterogeneous and depend on the context and quality of the child welfare system (Gross and Baron, 2022; Bald et al., 2022). This contested evidence underscores the importance of policies that promote timely exits from foster care to stable, permanent homes, which is the focus of our analysis.

Our findings demonstrate that international adoption has been a significant substitute for foster care adoption, particularly for the youngest children. Both our difference-in-differences and panel IV estimates show a robust, negative displacement effect. While international adoptions have declined from their mid-2000s peak, such trends can be volatile and may reappear in response to global events, such as the increased interest in adoption following the 2010 Haiti earthquake. Our results thus offer a wider view of the unintended consequences of policies that influence these flows. This economic perspective complements an active and complex debate on the welfare implications of international adoption, which involves deep sociological and psychological considerations for the child, the adoptive parents, and the birth parents (Selman, 2012), alongside serious concerns about corruption and child trafficking (Graff, 2008).

Our results have implications for government intervention in adoption markets. While the social benefits of foster adoptions are clear, an important question is whether current subsidies are effectively internalizing the externality. However, the U.S. currently employs dueling subsidies: international adoptions enjoy large dollar-for-dollar federal tax credits (approximately $15,000). To the extent that international adoptions respond to these tax subsidies, our results suggest that they displace adoptions out of foster care, countering the benefits of subsidies designed to promote foster adoptions (Buckles, 2013; Brehm, 2021). Policymakers should consider the costs associated with this displacement, while also weighing the potential external benefits of international adoptions for immigrant children and adoptive families.

Our results for Assisted Reproductive Technology (ART) also indicate a negative substitution effect, though the estimates are less precise than for international adoptions. The mechanism of this substitution is supported by work from Zaresani and Schmidt (2024), who find that more generous state-level insurance mandates for IVF treatment lead to increased utilization among older women, which is in turn associated with a lower rate of child adoption by this same demographic. Our cautious interpretation is consistent with a literature that has found mixed results; for instance, while Gumus and Lee (2012) find evidence of substitution in the opposite causal direction (higher adoptions to lower IVF use), others find no average effect of fertility insurance mandates on adoption rates (Cohen and Chen, 2010). Our imprecise estimates, alongside these varied findings, suggest the substitution effect is likely not uniform and highlight the need for further research into its heterogeneous nature, which depends on the preferences and institutional factors that shape family-formation decisions.

Taken together, our findings suggest that demand for children in U.S. foster care moves with the accessibility and cost of alternative family-formation routes. Although our empirical design seeks to identify the underlying substitution effects, we stop short of a full welfare assessment. Such an analysis would necessitate weighing the welfare of displaced foster children and their families against that of international adoptees – a complex task.

7 Conclusion

By jointly analyzing international adoptions and ART births, this paper shows that shocks to these external channels have causal, and often substantial, effects on the permanency outcomes for vulnerable children in the U.S. foster care system. The volatility in international adoptions during the 2000s displaced a significant number of foster children, and the continued rise of ART may apply ongoing pressure on their permanency outcomes.

While we identify key substitution effects, other demand-side drivers of adoption merit further investigation (Doepke and Kindermann, 2019). Our analysis underscores the importance of evaluating family-formation policies holistically. The findings that federal tax credits for international adoption can work at cross-purposes with subsidies for foster care, and that fertility insurance mandates may have spillover effects on the child welfare system, highlight the interconnectedness of these policy domains. Ultimately, careful research on these markets is crucial for designing policies that effectively serve the best interests of vulnerable children and their families.

References

Abramowitz, Joelle. 2014. "Turning back the ticking clock: the effect of increased affordability of assisted reproductive technology on women's marriage timing." *Journal of Population Economics*, 27(2): 603–633.

Alpert, Abby, David Powell, and Rosalie Liccardo Pacula. 2018. "Supply-Side Drug Policy in the Presence of Substitutes: Evidence from the Introduction of Abuse-Deterrent Opioids." *The Quarterly Journal of Economics*, 133(2): 821–862.

Argys, Laura M., and Brian Duncan. 2013. "Economic Incentives and Foster Child Adoption." *Demography*, 50(3): 933–954.

Bald, Anna, Joseph J. Doyle, Jr., Mical Gross, and Brian A. Jacob. 2022. "The Economics of Foster Care." *Journal of Economic Perspectives*, 36(2): 223–246.

Bartik, Timothy J. 1991. *Who Benefits from State and Local Economic Development Policies?* Kalamazoo, MI:W.E. Upjohn Institute for Employment Research.

Borusyak, Kirill, Peter Hull, and Xavier Jaravel. 2022. "Quasi-Experimental Shift-Share Research Designs." *The Review of Economic Studies*, 89(1): 181–213.

Brehm, Margaret E. 2021. "Taxes and adoptions from foster care: evidence from the Federal Adoption Tax Credit." *Journal of Human Resources*, 56(4): 1031–1072.

Buckles, Kasey S. 2013. "Adoption Subsidies and Placement Outcomes for Children in Foster Care." *Journal of Human Resources*, 48(3): 596–627.

Buckles, Kasey, William N Evans, and Ethan MJ Lieber. 2023. "The drug crisis and the living arrangements of children." *Journal of Health Economics*, 87: 102723.

Cohen, I. Glenn, and Daniel L. Chen. 2010. "Trading-off Reproductive Technology and Adoption: Does Subsidizing IVF Decrease Adoption Rates and Should It Matter?" *Minnesota Law Review*, 95: 485–553.

Cooper, Ryan, Joseph J. Doyle, Jr., and Andrés Hojman. 2024. "Effects of Enhanced Legal Aid in Child Welfare: Evidence from a Randomized Trial of Mi Abogado." *American Economic Review*, 114(7): 2310–48.

Dayal, Vikram, and Anand Murugesan. 2023. "Panel Data and Fixed Effects." In *Demystifying Causal Inference: Public Policy Applications with R*. 193–226. Springer.

Doepke, Matthias, and Fabian Kindermann. 2019. "Bargaining over babies: Theory, evidence, and policy implications." *American Economic Review*, 109(9): 3264–3306.

Doyle, Jr., Joseph J. 2007. "Child Protection and Child Outcomes: Measuring the Effects of Foster Care." *American Economic Review*, 97(5): 1583–1610.

Doyle, Jr., Joseph J. 2008. "Child Protection and Adult Crime: Using Investigator Assignment to Estimate Causal Effects of Foster Care." *Journal of Political Economy*, 116(4): 746–770.

Gateway, Child Welfare Information. 2022. "Trends in U.S. Adoptions: 2010-2019."

Goldsmith-Pinkham, Paul, Isaac Sorkin, and Henry Swift. 2020. "Bartik Instruments: What, When, Why, and How." *American Economic Review*, 110(8): 2586–2624.

Graff, Emil James. 2008. "The lie we love." *Foreign Policy*, , (169): 59.

Gross, Max, and E Jason Baron. 2022. "Temporary stays and persistent gains: The causal effects of foster care." *American Economic Journal: Applied Economics*, 14(2): 170–199.

Gumus, Gulcin, and Jong-Rong Lee. 2012. "Alternative Paths to Parenthood: IVF or Child Adoption?" *Economic Inquiry*, 50(3): 802–820.

Hansen, Mary Eschelbach. 2007. "Using Subsidies to Promote the Adoption of Children From Foster Care." *Journal of Family and Economic Issues*, 28(3): 377–393.

Hansen, Mary Eschelbach, and Bradley A. Hansen. 2006. "The Economics of Adoption of Children from Foster Care." *Child Welfare*, 85(3): 559–583.

Henry, Erin, and Dillon M. Walker. 2023. "The Effectiveness of Social Policy through the Tax System: Do State Adoption Tax Credits Increase Adoptions from Foster Care?" *Journal of the American Taxation Association*, 1–28.

Imai, Kosuke, and In Song Kim. 2019. "When should we use unit fixed effects regression models for causal inference with longitudinal data?" *American Journal of Political Science*, 63(2): 467–490.

Jaeger, David A, Joakim Ruist, and Jan Stuhler. 2018. "Shift-share instruments and dynamic adjustments: The case of immigration." *NBER Working Paper*, 24285.

Moriguchi, Chiaki. 2012. "The Evolution of Child Adoption in the United States, 1950-2010." Institute of Economic Research, Hitotsubashi University Discussion Paper A-582.

Rashid, Amani, and Glen R. Waddell. 2019. "Do Lawyers Increase the Rate of Adoption Among Foster Children?" Unpublished.

Rubin, David M, Evaline A Alessandrini, Chris Feudtner, David S Mandell, A Russell Localio, and Trevor Hadley. 2004. "Placement stability and mental health costs for children in foster care." *Pediatrics*, 113(5): 1336–1341.

Selman, Peter. 2012. "The global decline of intercountry adoption: What lies ahead?" *Social Policy and Society*, 11(3): 381–397.

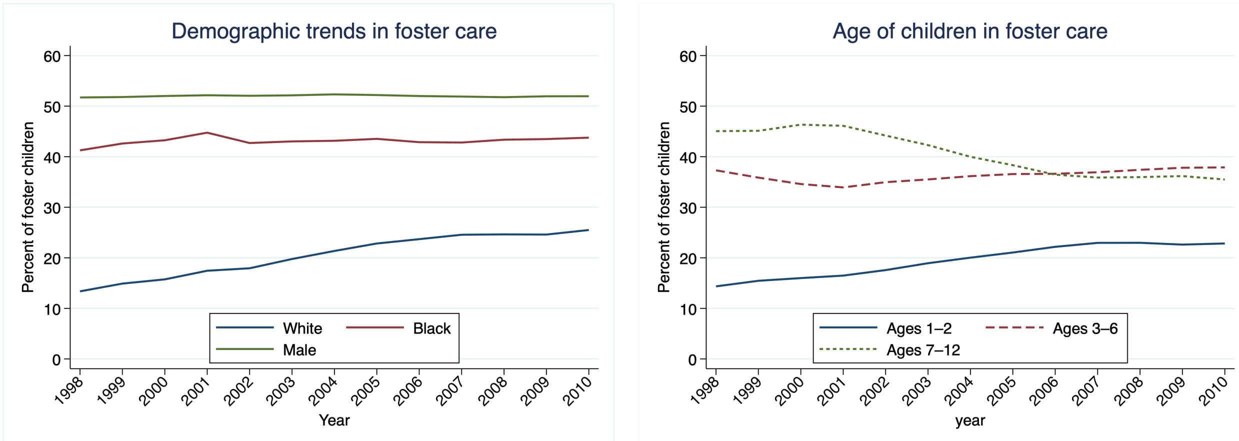
Shuman, Melissa, and Victor E. Flango. 2013. "Trends in U.S. Adoptions: 2000-2009." *Journal of Public Child Welfare*, 7(3): 329–349.

Skidmore, Mark, Gary M. Anderson, and Mark E. Eiswerth. 2016. "The Child Adoption Marketplace: Parental Preferences and Adoption Outcomes." *Public Finance Review*, 44(2): 163–196.

Swann, Christopher A., and M. Stephen Sylvester. 2006. "The Foster Care Crisis: What Causes Caseloads to Grow?" *Demography*, 43(2): 309–335.

Zaresani, Arezou, and Lucie Schmidt. 2024. “Unintended consequences of policy interventions: evidence from mandated health insurance coverage for IVF treatment.” *American Journal of Health Economics*, 10(3): 395–436.

Appendix



(a) Race (White/Black) and sex (Male) shares

(b) Age distribution (1–2, 3–6, 7–12)

Figure 3: Demography in foster care among children assigned adoption goal (1998 - 2010)

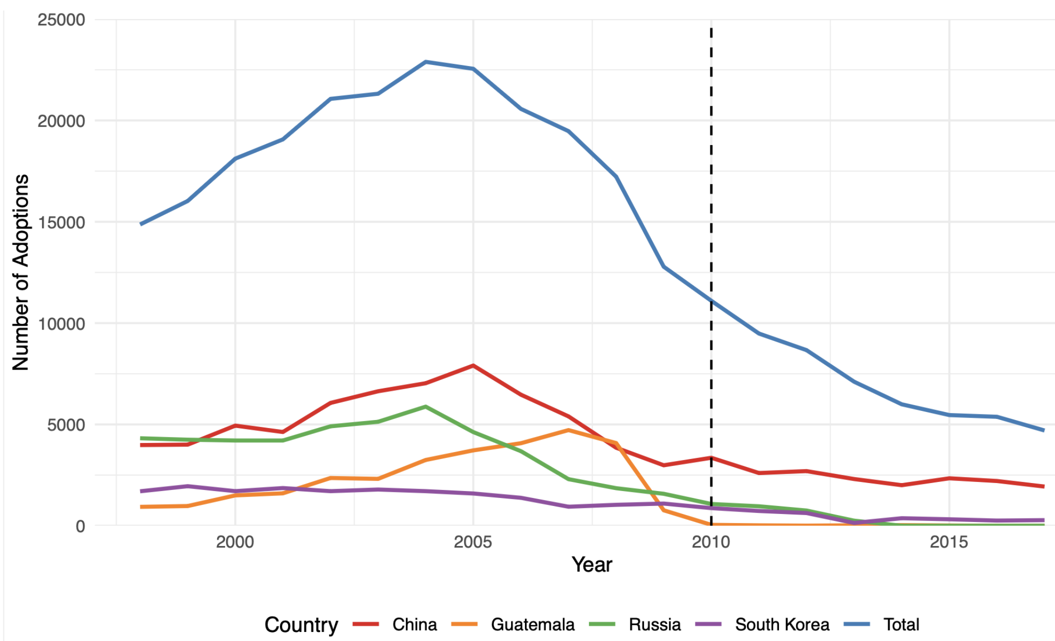


Figure 4: Trends in international adoptions from the major source countries

Table 5: Summary Statistics: State-Year Variables

Variable	Mean	Std. Dev.	N	Variable	Mean	Std. Dev.	N
Foster adoptions (all)	1057.55	1495.997	664	State fertility insurance	0.192	0.394	652
Foster adoptions ($\leq$ 4)	460.93	659.263	664	State adoption credits	0.189	0.392	652
Int'l adoptions	358.82	349.36	652	Triplicate state (dummy)	0.099	0.299	652
ART births	695.83	998.59	647	% College educated.	26.6	5.44	652
Int'l adoption rate $^+$	0.155	0.072	652	PCPI	33518.83	7473.81	652
ART birth rate $^+$	0.259	0.22	647	Median income	41426	8254.5	652
ART x Mandate IV $^+$	0.076594	0.22	641	Unemployment	166658.8	221503	652
Shift-share (10%) IV $^+$	0.105	0.07	652	Poverty (all age)	733290	898156	652

Rates marked $^+$ are per 1000 women of child-rearing age (15–44).

Table 6: Summary Statistics: Child Level (up to age 12; with case goal of adoption)

Variable	Mean	N	Variable	Mean	N
Adopted	0.297	1,710,117	Child disability	0.041	1,550,003
Waiting child	0.687	1,371,492	Behavioral problems	0.045	1,559,537
Parental rights terminated	0.617	1,371,492			
White	0.369	1,710,117	*Reasons for being in foster care*		
Amer. Indian	0.015	1,710,117	Neglected	0.643	1,559,692
Asian	0.005	1,710,117	Parents died	0.013	1,550,009
Afr. American	0.264	1,710,117	Parents in jail	0.066	1,550,011
Hawaiian/Pacific Islander	0.005	1,710,117	Inability to cope	0.222	1,559,593
Male	0.520	1,709,816	Abandonment	0.069	1,559,519
Parental drug abuse	0.275	1,559,593	Inadequate housing	0.136	1,549,815
Child drug abuse	0.041	1,549,874	Physically abused	0.171	1,559,632
Parental alcohol abuse	0.093	1,559,606	Sexually abused	0.054	1,559,633

Period 1998 - 2010, with repeat children. All variables are indicators (0–1); means are sample proportions.

Table 7: Attributes of U.S. Adopted Children and Adoption Costs by Source (% of Total)

Variable	Foster Care	Inter-national	Private Domestic	Variable	Foster Care	Inter-national	Private Domestic
Gender and race (2007)				**Age distribution at adoption (2007)**			
Male	57	33	51	Age less than 2	28	67	60
White	37	3	50	Age 2-5	42	25	21
Afr. Am.	35	9	25	Age 6-17	30	9	20
Asian	-	59	-				
Other (incl. hispanic)	28	19	25	**Child has special health care needs**			
				Children ages 0-5	39	10	25
Attributes of adoptive parents (2007)				Child w/ADD-ADHD	38	17	19
White	63	92	71	Child behavior problems	25	7	11
Afr. Am.	27	-	19				
Married	70	82	59	**U.S. adoption costs**			
No biological child	38	71	46	< \$1000	69	0	0
High school (or >) educ.	70	95	79	\$1-10,000	25	0	8
Related to child	23	-	41	\$10-20,000	6	1	23
Household income to poverty ratio				\$20-30,000	0	32	26
< 100% pov. line	16	-	17	\$30-40,000	0	51	25
100 - 400% pov. line	58	39	50	> \$40,000	0	16	18
> 400% pov. line	25	58	33				

Source: Adoption Factbook V (2012), Moriguchi (2012), *Adoptive Families* magazine

Table 8: Healthcare Costs Associated with In Vitro Fertilization (IVF)

Cost Category	**Description**	**Estimated Cost (USD)**
IVF Treatment	Cost for a single IVF cycle.	up to $15,000[a]
Total Healthcare Expenses (Maternal + Infant up to 1 year)[b]		
Singleton Birth	Total cost for one mother and one infant.	$27,454
Twin Birth	Total cost for one mother and two infants.	$114,697
Triplet (or more) Birth	Total cost for one mother and three or more infants.	$437,775

[a]Zaresani and Schmidt (2024), citing Lemos et al. (2013) among others, note a single cycle can cost up to 46% of a family's annual disposable income. Hamilton and McManus (2012) place out-of-pocket costs at $10,000-$15,000. States mandating insurance coverage of fertility treatments will drastically lower the costs.
[b]All birth-related costs are from Lemos et al. (2013) based on data from 2005-2010 (in 2010 US dollars). The primary driver of cost increases for multiple births is infant healthcare expenses, especially NICU stays.

Table 9: China (DiD) Pre-trends: Placebo shock in 2001

	Outcome: Foster Adoption Rate					
	Age 1	Age 2	Age 3	Age 4	Age 5	Age 6
3–year pre/post window						
Treated×Post	-0.007	-0.031	0.055	0.012	0.048	0.081
	(0.089)	(0.106)	(0.100)	(0.087)	(0.092)	(0.089)
Observations	17,696	23,617	24,006	22,986	21,807	20,472
4–year pre/post window						
Treated×Post	0.058	0.032	0.073.	0.010	0.052	0.089
	(0.074)	(0.091)	(0.082)	(0.071)	(0.070)	(0.066)
Observations	23,533	31,145	31,481	30,083	28,314	26,688

Notes: Entries are coefficients on Treated×Post with cluster-robust standard errors in parentheses (cluster: *state_year_id*). Significance: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

Table 10: Treated × Post estimates by country and pre/post window length

	Outcome: Foster Adoption Rate					
	Age 1	Age 2	Age 3	Age 4	Age 5	Age 6
Panel A: China						
3–year pre/post window						
Treated×Post	0.124**	0.140**	0.100*	0.066	0.084*	0.068*
	(0.050)	(0.062)	(0.054)	(0.048)	(0.047)	(0.039)
Observations	25,383	31,365	29,406	26,217	23,533	21,253
5–year pre/post window						
Treated×Post	0.115**	0.149***	0.131***	0.098***	0.114***	0.106***
	(0.044)	(0.049)	(0.045)	(0.037)	(0.037)	(0.036)
Observations	38,607	49,202	46,405	41,653	37,489	34,220
Panel B: Russia						
3–year pre/post window						
Treated×Post	0.132**	0.122*	0.160***	0.127***	0.127***	0.135***
	(0.062)	(0.063)	(0.050)	(0.044)	(0.046)	(0.046)
Observations	20,838	26,492	26,050	23,738	21,708	19,809
4–year pre/post window						
Treated×Post	0.136**	0.136**	0.173***	0.128***	0.145***	0.153***
	(0.056)	(0.058)	(0.047)	(0.041)	(0.042)	(0.041)
Observations	27,382	34,927	33,886	31,021	28,286	26,162
5–year pre/post window						
Treated×Post	0.114**	0.109*	0.145***	0.084*	0.101**	0.100**
	(0.052)	(0.057)	(0.050)	(0.044)	(0.046)	(0.049)
Observations	33,496	43,147	41,700	38,097	34,902	32,421

Notes: Entries are coefficients on Treated×Post with cluster-robust standard errors in parentheses (cluster: *state_year_id*). Significance: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

Table 11: IVProbit Estimates by Age Cohort (Ages 1–6), 1998–2010

	Age Cohort (years)					
	1	2	3	4	5	6
Panel A: International–adoption rate						
Coefficient	−7.549***	−10.47***	−6.266***	−6.425***	−1.344	−3.066
s.e. (robust)	(2.196)	(1.740)	(1.775)	(1.914)	(1.882)	(2.182)
Panel B: ART live-birth rate						
Coefficient	−8.120***	−5.974**	−6.180*	−7.770***	−6.695***	14.94***
s.e. (robust)	(1.560)	(1.936)	(2.403)	(2.052)	(2.018)	(0.692)

Notes: Standard errors are the robust s.e. shown directly beneath each coefficient in the source. $^{***}p < 0.001$, $^{**}p < 0.01$, $^{*}p < 0.05$.

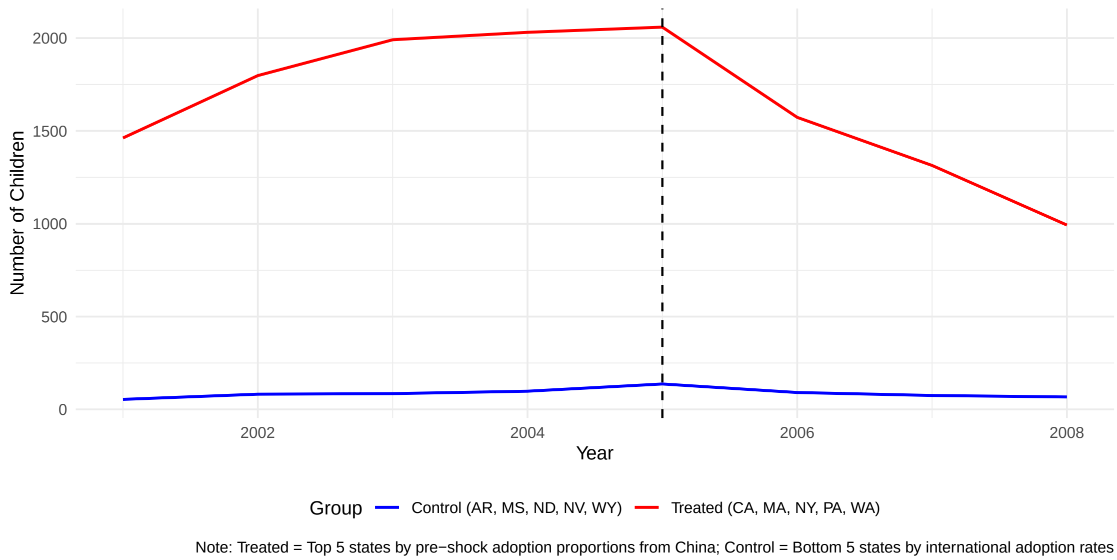


Figure 5: International Adoptions from China Only for High-Exposure (Treatment) vs. Low-Exposure (Control) States

Notes: The treatment group consists of the five states with the highest pre-shock adoption rates from China: California, New York, Pennsylvania, Massachusetts, and Washington. The control group consists of the five states with the lowest pre-shock rates. The vertical dashed line indicates the 2005 shock period.

7.1 A Simple Model of Adoption Markets

We consider a simple model of local adoption markets in order to frame our empirical hypotheses. Each family in a local population of potential child adopters faces four alternative strategies for securing a child: (1) ART (assisted reproductive technology), the biological alternative (subscript b)[9]; (2) a local private adoption (subscript p); (3) an international adoption (subscript i); or (4) a local foster adoption (subscript f).

Costs and Qualities. Costs of child acquisition under options (1)-(4) are C_k, $k \in \{b, p, i, f\}$. Each option is also associated with a measure of prospective (perceived) child quality PPCQ, q_k. All costs and PPCQ are exogenous except C_p (costs of private adoptions) and q_f (PPCQ quality for foster adoptions) which are market outcomes described below. We normalize PPCQ of a biological child to one, $q_b = 1$. In the adoption categories ($k \in \{p, i, f\}$), PPCQ q_k is an indicator of expected suitability of the match, assumed never superior to the biological alternative. For international adoptions, PPCQ q_i is determined (exogenously) in international

[9]Potential adopters have eschewed low-cost paths to own-conception. They are therefore candidates for either adoption or ART, and we do not consider the lesser (low cost) treatments for fertility enhancement.

markets. For local (independent) private adoptions that focus on easy placements of healthy children, PPCQ is plausibly higher than for international adoptions, $q_p > q_i$, as we assume in this Section.[10] In the Appendix, we consider the alternative of $q_p < q_i$.

Preferences. Each potential adopter is described by three preference parameters: overall preference for a child, K; preference for a biological vs. adopted child, $B \in [0, 1]$; and unit value of adopted child quality, $\alpha \in [0, \alpha_{max}]$. These preferences produce net benefits of the four options as follows (compared with the null alternative of no child, $U_0 = 0$):

ART: $U_b = KB - C_b$,

Adoption: $(k \in \{p, i, f\})$: $U_k = K(1 - B)\alpha q_k - C_k$.

Given its (B, α) preferences, each family chooses the one alternative that delivers highest utility. For simplicity, we assume that the overall preference parameter K is a common value that is sufficiently large to produce non-degenerate outcomes. The (B, α) preferences are heterogeneous with positive support $[0, 1]\text{x}[0, \alpha_{max}]$.

The Private (Independent) Adoption Market. The private adoption market is local and closed, with an exogenous and fixed supply of X_p^s .[11] The cost C_p is modeled as a price that clears the local market and includes non-monetary costs of the adoption process. For example, as demand for private adoptions rises, scarce (high PPCQ) children are rationed with higher costs of documenting parental suitability, longer queuing times, and higher non-pecuniary costs of rejection and reapplication, all reflected here in a higher C_p "price".

An alternative approach to modeling private adoptions is that the cost of private adoption C_p is fixed and both private and foster markets draw from a common pool of available children, distributed on the PPCQ (q) dimension. The private market skims the highest

[10]This premise reflects prevailing characterizations of the "baby shortage" in private adoption markets (e.g., Landes and Posner, 1978; Posner, 1992; Blackstone et al., 2004). Hansen (2007) writes: "When economists have written about adoption, they have primarily been interested in explaining why there are so few infants available for adoption through private agencies, lawyers, and facilitators, while there are so many prospective adoptive families who seek healthy infants." See also Medoff (1993) and Gennetian (1999) who study determinants of child supplies to private adoption markets.

[11] We can allow for a supply that is less than perfectly inelastic, $X_p^s(C_p)$. However, market-clearing variation in C_p is likely to be primarily non-monetary and not available to potential birth mothers (Landes and Posner, 1978), motivating the inelastic supply assumed here. A potential alternative specification for private adoptions treats them as exogenous to the model, viz, a segmented market in which costs are lower than international counterparts, $C_p < C_i$, and PPCQ is higher, $q_p > q_i$; in such a market, private adoptions dominate international adoptions and are rationed to qualifying "high quality" parents.

PPCQ children, and foster adoptions come from the highest PPCQ residual children, those not adopted in the private market.[12] An equilibrium is described by average PPCQ qualities in private and foster adoption markets, with $q_p > q_f$.

Foster Adoptions. Let N be the number of foster children available for adoption, those for whom authorities have ruled out reunification with parents. Within this pool, PPCQ q is distributed with density/distribution $f(q)/F(q)$ on the support, $[\underline{q}, \overline{q}], 0 < \underline{q} < \overline{q} \leq 1$. Potential adopters observe only the average PPCQ of all prospective foster adoptees, q_f, when deciding whether to pursue a foster adoption strategy. Whether by strategic prioritization of children waiting to be adopted or by the process of child-adopter matching, we assume that the highest PPCQ children are adopted first. Hence, q_f is determined by the pool of foster adoptees in the upper tail of the PPCQ distribution.

Once potential adopters observe q_f (and attributes of the alternatives), they each commit to one strategy. Those committing to foster adoption are matched to a child in the pool of (highest PPCQ) foster children, with actual PPCQ of the match randomly assigned from within this pool. For simplicity we assume that foster demanders do not observe true matched child PPCQ before adoption.[13]

Because foster adoptions skim the top PPCQ children, $q > \underline{q}_f$, average PPCQ in the foster adoption pool, q_f, is:

$$q_f = \int_{\underline{q}_f}^{\bar{q}} \frac{q f(q)}{(1 - F(\underline{q}_f))} \, dq \leftrightarrow \underline{q}_f = \underline{q}_f(q_f) \tag{3}$$

[12] Baccara et al. (2012) study matching of parents and children in private adoption markets. Children not matched in the private market enter foster care (Baccara et al., 2012, p. 12).

[13] If foster demanders observe true child PPCQ before adoption, they can accept or reject the matched child and may reject if q is "too low." However, because demanders have committed to the foster adoption, "rejecters" are out of the child market for the time period of the model. This twist to the model has two symptoms: (1) due to rejecters, actual adoptions are fewer than the number of foster adoption demanders, and (2) the utility of the foster adoption strategy, U_f, is higher than described in this Section due to the option value of rejection. Despite these symptoms, key qualitative implications of the model persist: U_f rises with q_f and there is a monotonic relationship between the number of adoption demanders and actual foster adoption numbers (and associated adoption probabilities). Moreover, implications of the analysis for our empirical approach (netting out of private adoptions and use of substitute quantities to proxy for costs) withstand the twist.

producing the total foster adoption supply,

$$X_f^s(q_f) = N(1 - F(\underline{q}_f(q_f))) \leftrightarrow \frac{\partial X_f^s}{\partial q_f} = -Nf(\underline{q}_f)(\partial \underline{q}_f / \partial q_f) < 0 \tag{4}$$

and probability of foster adoption $H = (1 - F(\underline{q}_f))$. A higher foster adoption supply requires ratcheting down the quality distribution, depleting average q_f.

Demand for Adoption Alternatives. We begin with how potential adopters choose between the adoption alternatives p, i, f and then follow with the addition of the biological/ART alternative. Throughout, we assume the following:

Assumptions: (A1) In equilibrium, there is positive local demand for all strategies to secure a child, $k \in \{b, p, i, f\}$. (A2) Relative to foster adoption, (i) international adoptions are more costly, $C_i > C_f$, and (ii) international and private adoptions have high PPCQ qualities, $q_i > E(q) = \int_{\underline{q}}^{\bar{q}} qf(q)dq$ and $q_p \geq \bar{q}$.

Assumption (A1) focuses us on the relevant cases of positive demands. Recalling the Table 1 statistics on special needs, PPCQ is "high" for both international and private adoptions relative to the foster counterpart, particularly for the lowest ages that dominate the substitute categories. Assumption (A2) reflects these realities. Together, the two assumptions imply an ordering of cost and PPCQ:

Observation 1: With $q_p > q_i$, $C_p > C_i > C_f$ and $q_p > q_i > q_f$ in equilibrium.[14]

Mapping choices in the (B, α) parameter space, the set of agents indifferent between two respective alternatives can be defined as follows:

$$\alpha_{kj}(B) : K(1 - B)\alpha q_k - C_k = K(1 - B)\alpha q_j - C_j,\ kj \in \{pi, if, pf\} \tag{5}$$

$$\alpha_{f0}(B) : K(1 - B)\alpha q_f - C_f = U_0 = 0 \tag{6}$$

[14] If $C_p \leq C_i$ (and $q_p > q_i$), then private adoptions dominate international, contrary to Assumption (A1); hence, $C_p > C_i$. Similarly, if $q_i \leq q_f$ (and $C_i > C_f$ by Assumption (A2)), then foster adoptions dominate international; hence (again by (A1)), $q_i > q_f$.

Eq. (3) defines indifference between private and international adoptions ($kj = pi$), international and foster adoptions ($kj = if$), and private and foster ($kj = pf$). Eq. (4) defines indifference between foster adoption and the null (no child) alternative. The indifference mappings have the following properties:

$$\frac{d\alpha_A(B)}{dB} = \frac{\alpha_A}{(1-B)}, \quad A \in \{pi, if, pf, f0\} \tag{7}$$

$$\alpha_{max} > \alpha_{pi}(0) > \alpha_{if}(0) > max(0, \alpha_{f0}(0)) \tag{8}$$

With $q_p > q_i$, eq. (6) is necessary for positive adoption demands ($k \in \{p, i, f\}$).

Figure 6 graphs the indifference mappings (left panel), indicating respective preference for the four alternatives $p, i, f, 0 =$ null. Note the hierarchy of preferences: Private adoptions are made by agents in the top-most space of preferences; international adoptions are made by agents in the next highest space, followed below by foster adoptions and, finally, in the lowest quadrant, agents who prefer not to adopt (those with low α and/or high B).[15]

The ART Alternative. The final component of demand is ART. The minimum B preference that makes ART preferable to no child is $\underline{B} \equiv C_b/K$, where $\underline{B} < 1$ by Assumption (A1). The set of agents indifferent between ART and adoption alternative $k \in \{p, i, f\}$ is:

$$\alpha_{bk}(B) : KB - C_b = K(1-B)\alpha q_k - C_k \tag{9}$$

where

$\frac{d\alpha_{bk}}{dB} = \frac{1+\alpha q_k}{(1-B)q_k} > \frac{\alpha}{(1-B)} > 0, \quad \frac{\partial}{\partial q_k}\left[\frac{d\alpha_{bk}}{dB}\right] = -\frac{1}{(1-B)q_k^2} < 0,\ \alpha_{bf}(\underline{B}) = \alpha_{f0}(\underline{B})$

The (B, α) pair that yields zero net benefit of both ART and foster adoption, $(\underline{B}, \alpha_{f0}(\underline{B}))$, also leaves the agent indifferent between the two (ART and foster) alternatives. Adding these

[15]For the alternate case of $q_i > q_p$, international and private adoptions swap places in the Figure 5 representation of demands (see on-line Appendix). In this case, substitution between international and foster adoptions (due to a change in C_i) operates through the private market; a rise in C_i leads to an equilibrium rise in C_p, which raises demand for foster adoptions.

constructs to Figure 5 produces the child demands:

$$X_k = X_k^D(C_p, C_i, C_b, C_f, q_f) = \text{ demand for strategy } k \in \{b, p, i, f\}. \tag{10}$$

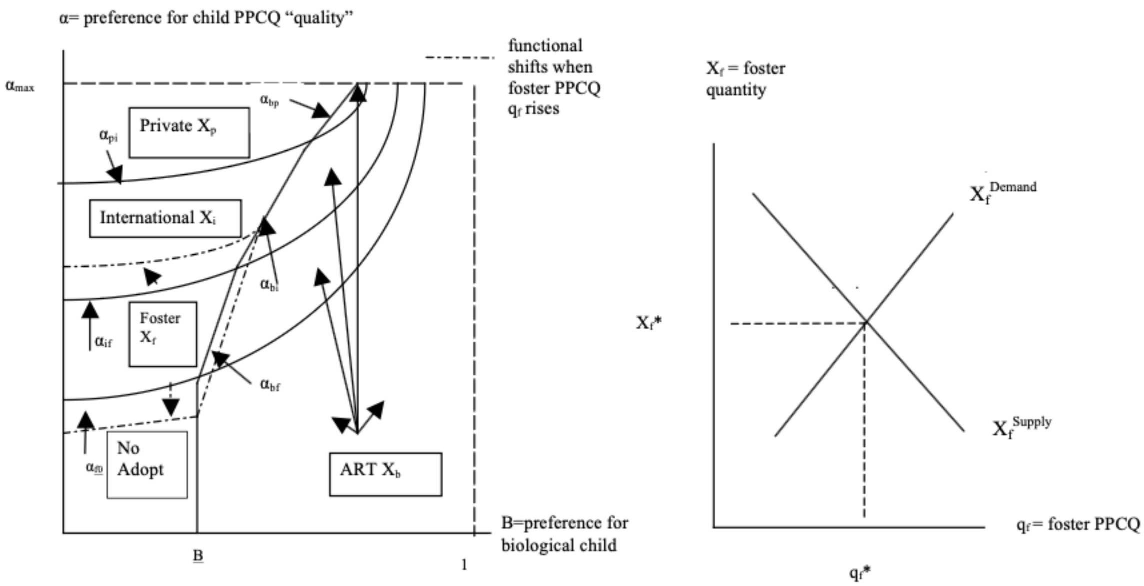


Figure 6: Indifference mappings (left); Equilibrium in the foster adoption market (right)

The model is closed by an equilibrium price in the private market, C_p, and an equilibrium quality in the foster market, q_f, that equate supplies and demands:

$$C_p : X_p^D(C_p, C_i, C_b) = X_p^S \tag{11}$$

$$q_f : X_f^D(C_i, C_b, C_f, q_f) = X_f^S(q_f) \tag{12}$$

Our primary interest is the foster market equilibrium, equation (10), depicted in (right panel) Figure 6 . On the demand side (and referring to Figure 6), a higher quality of foster matches has three complementary effects: (1) the α_{f0} margin for positive foster net benefits shifts down; (2) the α_{if} margin for preference of foster vs. international adoptions shifts up, and (3) the α_{bf} margin for preference of foster vs. ART children shifts out. The three shifts are illustrated by dotted lines in Figure 5. All lead to increased demand for foster adoption, captured by the upsloping demand function in Figure 6. Conversely, on the supply side, a higher supply of foster adoptees requires ratcheting down the PPCQ spectrum for promising adoptive matches in the foster system; as a result, the average q_f falls – giving us the down-sloping foster supply function in Figure 6.

For our empirical analysis, the key comparative statics are for the costs of international adoption and ART, C_i and C_b. Higher costs shift the foster demand function X_f^D out (Figure 6), leading to a higher equilibrium level of adoptions (X_f^*) and a lower level of PPCQ, q_f^* . In terms of the primitive drivers of demand in Figure 5, a higher C_i shifts the α_{if} curve up, increasing foster demand. A higher C_b shifts the $\underline{B}$ and α_{bj} margins to the right, also increasing foster demand.

Although qualitative effects of the two substitute categories (international and ART) are similar, quantitative effects can be very different. Suppose, for example, that the distribution of B preferences is bimodal with almost all families having either low values of B (families who prefer to adopt) or high values of B (families who prefer a biological child). In this case, a change in costs of ART (C_b) will have negligible effect on demand for foster children as there are almost no families in the range of B values for which B and α_{bf} margins shift. However, a change in costs of international adoption (C_i) will affect agents in the low-B-value range, having a substantial effect on foster demand.

Summarizing our model predictions for the empirical analysis:

Observation 2 (Main Hypothesis). A higher cost of international adoptions (C_i) or ART (C_b) increases adoptions out of foster care, $\partial X_f^*/\partial C_k > 0$ for $k \in \{b, i\}$.

In our empirical work, costs are unobservable and we use observable proxies for costs, namely, quantities. The following observation implies invertible relationships between the equilibrium quantities (X_i and X_b) and respective costs (C_i and C_b). We exploit these relationships to infer substitution effects ($\partial X_f^*/\partial C_i$ and $\partial X_f^*/\partial C_b$) from estimated effects of normalized quantities (X_i and X_b, normalized by a relevant population of potential adopters) on foster adoption probabilities.

Observation 3. (A) Equilibrium levels of international adoptions and ART demand are negatively related to own cost: $dX_i/dC_i < 0$ and $dX_b/dC_b < 0$. (B) Costs C_i and C_b have the same signs of effect on the equilibrium foster demand X_f^* and on the equilibrium probability of foster adoption:[16]

$$\frac{dH}{dC_k} = -f(\underline{q}_f(q_f^*))\,\mathrm{x}\,\frac{d\underline{q}_f}{dq_f}\,\mathrm{x}\,\frac{\partial X_f^*/\partial C_k}{\partial X_f^S/\partial q_f} \stackrel{\mathrm{s}}{=} \frac{\partial X_f^*}{\partial C_k}\ , \quad k \in \{i, b\} \tag{13}$$

where H = probability of adoption from foster care $= 1 - F(\underline{q}_f(q_f^*))$.

[16] In equation (11), we substitute for $\partial q_f^*/\partial C_k$ from the equilibrium relationship, $X_f^* = X_f^S(q_f^*)$, and the sign equality follows from $d\underline{q}_f/dq_f > 0$ and $\partial X_f^S/\partial q_f < 0$.