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Working Paper

Comment on "Diffusion of Reproductive Health Behavior through International Migration: Effects on Origin-Country Fertility"

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No. 327

DISCUSSION PAPER SERIES

Comment on “Diffusion of Reproductive Health Behavior through International Migration: Effects on Origin-Country Fertility”

Simone Bertoli

Jules Gazeaud

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I4R DISCUSSION PAPER SERIES

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Comment on “Diffusion of Reproductive Health Behavior through International Migration: Effects on Origin-Country Fertility”

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Comment on “Diffusion of Reproductive Health Behavior through International Migration: Effects on Origin-Country Fertility”*

Simone Bertoli, Jules Gazeaud

June 4, 2026

Abstract

Godlonton and Theoharides (2025) argue that Filipino migrants’ exposure to more liberal reproductive health policies reduces origin-community fertility. This comment shows that the differential effect they present is insufficient to conclude that more liberal policies significantly reduce fertility at origin—on average or in any single province. These claims require estimates of average or province-specific effects, but neither the article nor the replication files include information to recover them. The Authors declined our requests to share portions of the analysis data at the province-year level or information on a key coefficient estimated by their code but left unreported. We overcome this otherwise binding data constraint using information from their working paper. We show that the average effect is far from significant ($|t\text{-stat}| \leq 0.86$), and that the province-specific effect may be negative and significant for as few as one out of 84 provinces. This occurs even though the differential effect is negative and highly statistically significant, as in Godlonton and Theoharides (2025).

*Godlonton, Susan and Caroline Theoharides, “Diffusion of Reproductive Health Behavior through International Migration: Effects on Origin-Country Fertility,” *American Economic Review*, October 2025, 115 (10), 3597–3637; Bertoli: Université Clermont Auvergne, CNRS, IRD, CERDI (email: simone.bertoli@uca.fr). Gazeaud: Université Clermont Auvergne, CNRS, IRD, CERDI (email: jules.gazeaud@uca.fr). We are grateful to Susan Godlonton and Caroline Theoharides for having taken the time to reply to our requests; the code and data underlying this comment are available at: <https://y.uca.fr/QkYw3tfm>; the usual disclaimers apply.

1 Introduction

Godlonton and Theoharides (2025), hereafter GT or the original article, study how norms and ideas diffuse across countries through international migration. GT focus on temporary Filipino migrants and examine whether changes in destination-country reproductive health policies affect fertility and related outcomes in the Philippines.

GT use rich data to construct a shift-share variable exploiting temporal variation in reproductive health policies across destination countries (the shifters) along with spatial variation in the historical distribution of migrants from each province across destinations (the shares). Specifically, GT extend a country-year-level dataset of reproductive health policies over 1992–2013 produced by Finlay et al. (2012), which they combine with information on the distribution of migrants from 84 Filipino provinces across destination countries in 1992. GT refer to this shift-share variable as the “weighted liberalization index” or $WeightedLibIndex_{p,t}$. Because GT expect effects to be stronger in provinces with more migrants, they interact the weighted liberalization index with the baseline migration rate $MigRate_{p,t=0}$ of province p . They then regress the fertility rate—and other outcomes of interest such as contraceptive use, infant mortality, and maternal mortality—on this interaction, the uninteracted weighted liberalization index, and various fixed effects, including province dummies which absorb the uninteracted baseline migration rate:¹

$$Y_{pt} = \beta_0 + \beta_1 WeightedLibIndex_{p,t-2} + \beta_2(MigRate_{p,t=0} \times WeightedLibIndex_{p,t-2}) + \delta_p + \tau_t + \delta_p \times time_t + \epsilon_{pt}$$

GT indicate that β_2 , i.e., the coefficient of the interaction term, is “[their] primary coefficient of interest” (p. 3609), and infer from its sign and statistical significance that “exposure” to more liberal policies significantly reduces fertility at origin.^{2,3}

In this comment, we contend that evidence on β_2 does *not* suffice to conclude that

¹See Eqs. (1) and (2) in Section 2 of this comment for more details.

²For instance, the Abstract states that “[m]igrant exposure to more liberalized reproductive health policies reduces origin-community fertility”, and similar statements can be found throughout the original article. Introduction: “We find that provinces exposed to more liberalized reproductive health policies experienced a decline in the fertility rate” (p. 3599). Main text: “Exposure to more liberalized policies has a statistically significant effect on fertility” (p. 3617). Conclusion: “In this paper, we examine the effect of exposure to more liberalized reproductive health policies on fertility decisions in the Philippines. We find that fertility declines in response to that” (p. 3633).

³GT use the expression “policy exposure” to refer both to the weighted liberalization index (e.g., at pp. 3608 and 3610) and to its interaction with the baseline migration rate (e.g., at pp. 3609 and 3614). We refrain from using this expression because it is unclear what this refers to.

more liberal policies lead to a significant reduction in fertility at origin, or that such a reduction is significant for a subset of provinces with more migrants. The province-specific effect of a one standard deviation increase in the weighted liberalization index is equal to $\hat{\beta}_1 + \hat{\beta}_2 \times MigRate_{p,t=0}$, and the average effect is equal to this expression evaluated at the mean baseline migration rate. These effects are not reported in the original article, and the estimate of β_1 is never disclosed, including in the Supplemental Appendix and in the replication files.

What can be inferred from estimates of β_2 alone? For each specification, GT focus the discussion on the quantity $\hat{\beta}_2 \times IQR$, where IQR is the interquartile range of the distribution of the baseline migration rate, which stands at 0.25 percentage points. We show that the quantity $\hat{\beta}_2 \times IQR$ corresponds to the *differential* effect of a one standard deviation increase in the weighted liberalization index between any two provinces for which the baseline migration rates differ by 0.25 percentage points. The statistical significance of this quantity clearly coincides with that of $\hat{\beta}_2$. Crucially, however, this quantity is not informative in itself about the sign, the magnitude, and the statistical significance of the effect of more liberalized policies on fertility—on average or in any single province. For instance, the province-specific effect on fertility may be positive in all provinces if $\hat{\beta}_1$ is positive and sufficiently large to offset the negative effect of $\hat{\beta}_2$. Conversely, the average effect and the province-specific effect may be more negative than the differential effect if $\hat{\beta}_1 < 0$.⁴ Assessing these effects requires information on $\hat{\beta}_1$, its standard error, $Cov(\hat{\beta}_1, \hat{\beta}_2)$, and the baseline migration rate of the 84 Filipino provinces in the sample—none of which is reported in GT or in the replication files.

When GT's abstract argues that “[m]igrant exposure to more liberalized reproductive health policies reduces origin-community fertility”, this should be interpreted narrowly as evidence of the significant role of migration intensity in shaping the change in fertility across provinces induced by a liberalization of reproductive health policies at destination. Such an interpretation fails to unveil the richness of the evidence produced by the specification estimated in GT, which allows identifying the lowest migration intensity leading to a significant effect on fertility and related outcomes. Furthermore, the Authors go beyond this narrow interpretation in at least two instances: (i) in their presentation of the results in a non-technical summary, and (ii) in the original article when they draw implications from the estimates. With respect to point (i), the Authors state that “[l]ess restrictive reproductive health policy abroad reduces fertility at home” and that “fertility falls in response to larger reductions in legal restrictions regarding reproductive health in migrants’ desti-

⁴These remarks also apply to the other outcomes of interest in GT, e.g., infant and maternal mortality.

nations” (Godlonton and Theoharides, 2026), but these conclusions cannot be drawn from the differential effect they estimate. As far as point (ii) is concerned, GT interpret their differential effect as an average effect to conclude that “effect sizes must encompass a fertility response by both women in migrant and nonmigrant households” (p. 3619).⁵

To understand whether more liberal policies lead to a significant reduction in fertility at origin, and to quantify the size of this reduction, we sought to derive the average and province-specific effects in GT.⁶ However, as mentioned above, one challenge is that the replication package includes neither the analysis data nor the estimation results. We attempted to overcome this constraint in two ways. First, we contacted the Authors to request limited portions of the analysis data at the province-year level, but they responded that they cannot share these data beyond the AEA replication team due to two MOUs and a data agreement they had signed. Second, we asked the Authors if they could instead share estimates from four key regressions included in the original article, but they declined our request. We considered the option of conducting a “deep” replication, starting from the raw data. The fact that the data must be obtained from several providers, one of which was visited in person in the Philippines (as mentioned in the readme file of the replication package), suggests that such a deep replication would not be feasible with reasonable resources.⁷

Even though we face binding constraints with respect to the original article, an earlier working paper version (Godlonton and Theoharides, 2022, GT22 hereafter) allows us to provide evidence of the extent to which the differential effect cannot be reliably used to infer the province-specific or the average effect. GT22 estimate a specification at the municipality-year level that is closely related to the equation in GT, but, contrary to GT, both $\hat{\beta}_1$ and

⁵GT rely on the country-wide share of women aged 15 to 49 living in households with a migrant in the 1990 census, i.e. 3.18 percent, and argue that a 10 percent reduction in fertility for women in these migrant households “would explain about a third of the overall [0.96 percent] reduction in fertility” ($0.0318 \times 0.10 \approx 0.0096/3$). This insight motivates the discussion in Section VID (Diffusion to Nonmigrants). However, the computation should have been done using the difference in the share of women living in migrant households for provinces with a baseline migration rate corresponding to the seventy-fifth and to the twenty-fifth percentile of the distribution (or, more broadly, for provinces whose baseline migration rate differ by IQR), and not with the country-wide share of women living in migrant households.

⁶We pre-registered our analyses on February 11, 2026, after the original article was published but before the replication package was available online, see <https://osf.io/5vj3c/overview>. Appendix A describes how this comment relates to the tests and discussion we pre-registered.

⁷In addition, the many-analyst study by Huntington-Klein et al. (2025) shows that variations in data cleaning decisions across research teams can lead to variations in estimated effects, suggesting that our ability to replicate the results in the original article is ex ante uncertain. The replication package does include the cleaning do-files used by the Authors, but we are uncertain we could obtain the same versions of the raw data, as in our experience data providers such as national statistics offices frequently update information systems, making old versions of administrative datasets difficult or impossible to retrieve in previous formats.

$\hat{\beta}_2$ are reported, together with information on the distribution of the baseline migration rate across provinces, and on the quintile to which each province belongs to. We show that the province-specific effect of a one standard deviation increase in the weighted liberalization index on fertility in the original article may be negative and significant for as few as one province and that the average effect on fertility in GT22 is far from being statistically different from zero ($|t\text{-stat}| \leq 0.86$), in contrast with the differential effect ($|t\text{-stat}| \geq 3.07$).

While this evidence does *not* necessarily imply that the average effect in the original article is also insignificant, it clearly reveals that the sign, size and significance of the differential effect are not informative about the average effect, or about the share of provinces for which the effect on fertility of a one standard deviation increase in the weighted liberalization index is negative and significant. One might reasonably expect limited or no effects in provinces with low migration intensity, as cultural remittances induced by international migration are likely to arise only if migration intensity is sufficiently high. A non-significant average effect, or the presence of non-significant point estimates for some provinces, would therefore not necessarily be concerning for the hypothesis that Filipino migrants transmit changes in reproductive health policies from destination to origin communities. However, identifying the threshold above which effects turn significant appears crucial to understand the extent of this transmission.

The remainder of this comment is organized as follows: Section 2 presents the province-specific, average and differential effects; Section 3 derives the size and significance of the province-specific and of the average effects drawing on the information in GT22; Section 4 concludes.

2 Province-specific, average and differential effect

The weighted liberalization index is defined as follows in Eq. (1) in GT:

$$WeightedLibIndex_{p,t} = \sum_d (MigShare_{pd0} \times LibIndex_{dt}) \quad (1)$$

where $MigShare_{pd0}$ is the share of migrants from province p in country d in 1992 ($t = 0$), and $LibIndex_{dt}$ is the value of the liberalization index in country d and year t . In the econometric analysis, $WeightedLibIndex_{p,t}$ is lagged by two years and standardized to have mean zero and a unitary standard deviation. GT interact $WeightedLibIndex_{p,t-2}$ with the

baseline migration rate $MigRate_{p,t=0}$ and estimate the following specification (corresponding to Eq. (2) in GT):

$$Y_{pt} = \beta_0 + \beta_1 WeightedLibIndex_{p,t-2} + \beta_2 (MigRate_{p,t=0} \times WeightedLibIndex_{p,t-2}) + \delta_p + \tau_t + \delta_p \times time_t + \epsilon_{pt} \quad (2)$$

where δ_p and τ_t are province and time fixed effects respectively, and $\delta_p \times time_t$ are linear province-specific time trends. The dependent variable Y_{pt} describes different outcomes throughout the original article, most notably the fertility rate (derived from population censuses and the Philippine Demographic and Health Survey, PDHS), contraceptive use (PDHS), and infant and maternal mortality (Vital Statistics from the Philippines).

2.1 Province-specific effect

From Eq. (2), we can easily derive ΔY_p , the change in the dependent variable induced by a one standard deviation increase in the weighted liberalization index:

$$\Delta Y_p = \hat{\beta}_1 + \hat{\beta}_2 \times MigRate_{p,t=0} \quad (3)$$

ΔY_p is province-specific, as it depends on the baseline migration rate $MigRate_{p,t=0}$, and it is a function of β_1 and β_2 from Eq. (2). If $\hat{\beta}_1 \times \hat{\beta}_2 < 0$, then ΔY_p changes sign when $MigRate_{p,t=0} = M^{l_1} \equiv |\hat{\beta}_1 / \hat{\beta}_2|$. For example, if $\hat{\beta}_1 > 0$ and $\hat{\beta}_2 < 0$, then $\Delta Y_p < 0$ only for a province p such that $MigRate_{p,t=0} > M^{l_1}$. The standard error $\hat{\sigma}_{\Delta Y_p}$ of ΔY_p in Eq. (3) is equal to:

$$\hat{\sigma}_{\Delta Y_p} = \left(\hat{\sigma}_{\beta_1}^2 + (MigRate_{p,t=0})^2 \times \hat{\sigma}_{\beta_2}^2 + 2MigRate_{p,t=0} \times \text{Cov}(\hat{\beta}_1, \hat{\beta}_2) \right)^{1/2} \quad (4)$$

where $\hat{\sigma}_{\beta_1}$ and $\hat{\sigma}_{\beta_2}$ denote respectively the standard errors of $\hat{\beta}_1$ and $\hat{\beta}_2$, and $\text{Cov}(\hat{\beta}_1, \hat{\beta}_2)$ is the covariance between $\hat{\beta}_1$ and $\hat{\beta}_2$. The null hypothesis that the province-specific effect is equal to zero can be tested by computing the t -statistic $\Delta Y_p / \hat{\sigma}_{\Delta Y_p}$ from Eqs. (3)-(4). It is straightforward to demonstrate that when $\hat{\beta}_1 > 0$ and $\hat{\beta}_2 < 0$, then the t -statistic $\Delta Y_p / \hat{\sigma}_{\Delta Y_p}$ from Eqs. (3)-(4) implicitly defines a threshold $M^{l_2}(t) > M^{l_1}$ such that ΔY_p is negative and significant only for a province p such that $MigRate_{p,t=0} > M^{l_2}(t)$, where t is the absolute value of the t -statistic above which a coefficient is considered as statistically significant. More precisely, $M^{l_2}(t)$ corresponds to the right root of the equation $\Delta Y_p / \hat{\sigma}_{\Delta Y_p} = t$, which is equal

to:⁸

$$M^{l_2}(t) = \frac{-(\hat{\beta}_1\hat{\beta}_2 - t^2\text{Cov}(\hat{\beta}_1, \hat{\beta}_2)) + [(\hat{\beta}_1\hat{\beta}_2 - t^2\text{Cov}(\hat{\beta}_1, \hat{\beta}_2))^2 - (\hat{\beta}_2^2 - t^2\hat{\sigma}_{\hat{\beta}_2}^2)(\hat{\beta}_1^2 - t^2\hat{\sigma}_{\hat{\beta}_1}^2)]^{1/2}}{(\hat{\beta}_2^2 - t^2\hat{\sigma}_{\hat{\beta}_2}^2)} \quad (5)$$

2.2 Average effect

As the province-specific effect in Eq. (3) is an affine function of the baseline migration rate, its average value, which we denote $\bar{\Delta Y}$, corresponds to ΔY_p evaluated at the mean baseline migration rate $\bar{M} = 0.0022 = 0.22$ percent (this mean is shown in Panel D – Table A2 in the Supplemental Appendix of GT), that is:

$$\bar{\Delta Y} = \Delta Y_p(\bar{M}) = \hat{\beta}_1 + \hat{\beta}_2 \times 0.0022$$

Its statistical significance can be assessed by evaluating Eqs. (3)-(4) for $MigRate_{p,t=0} = \bar{M}$.

2.3 Differential effect

The main tables in GT include a row labeled “Coefficient $\times IQR$ ”, where “Coefficient” is $\hat{\beta}_2$, and IQR denotes the interquartile range of the baseline migration rate (i.e., 0.0025, or 0.25 percent). The statistical significance of this quantity coincides with the statistical significance of $\hat{\beta}_2$. For instance, for Col. (1) in Panel A of Table 4, where the dependent variable is the fertility rate computed from the census and $\hat{\beta}_2 = -0.516$ (s.e. 0.187), “Coefficient $\times IQR$ ” corresponds to $-0.516 \times 0.0025 = -0.0013$ (–0.13 percentage points, or a 0.96 percent reduction with respect to the average yearly fertility rate).

The quantity “Coefficient $\times IQR$ ” describes the differential effect of a one standard deviation increase in the weighted liberalization index between provinces whose baseline migration rates differ by the IQR . Formally, “Coefficient $\times IQR$ ” corresponds to:

$$\hat{\beta}_2 \times IQR = \Delta Y_p(MigRate_{p,t=0} + IQR) - \Delta Y_p(MigRate_{p,t=0}),$$

Importantly, this quantity just rescales the effect of $\hat{\beta}_2$, and it does not provide information about the average effect of more liberalized policies $\bar{\Delta Y}$, nor does it speak about the heterogeneity of the province-specific effect ΔY_p along the baseline migration rate.

⁸If $\hat{\beta}_1$ is positive but not significant, i.e., $\hat{\beta}_1/\hat{\sigma}_{\hat{\beta}_1} \leq t$, then the equation $\Delta Y_p/\hat{\sigma}_{\Delta Y_p} = t$ admits only one positive root; this, in turn, implies that ΔY_p is positive when $MigRate_{p,t=0} < M^{l_1}$, but it is not significantly different from zero.

3 Evidence on the province-specific and average effect

As described in the Introduction, we could not derive the average and province-specific effects in GT due to binding data constraints. Yet, an earlier working paper version of GT (Godlonton and Theoharides, 2022, hereafter GT22) allows to provide clear evidence that the differential effect is not informative about the size and significance of the province-specific and average effects.

GT22 provide estimates of β_1 and more information on the distribution of the province-specific baseline migration rate.⁹ They estimate a specification at the municipality-year level that is closely related to Eq. (2) in GT. Specifically, Col. (1) in Panels A and B of Table 4 reports estimates of a regression in which the fertility rate (measured from the census and the PDHS respectively) is regressed on the baseline migration rate, the weighted liberalization index, and their interaction, as in GT. These two specifications report estimates for β_1 and β_2 with opposite signs, i.e., $\hat{\beta}_1 > 0$ and $\hat{\beta}_2 < 0$.¹⁰

Based on these estimates and on the different values of the baseline migration rate reported in GT and GT22, we show in Figure 1 the effect of a one standard deviation increase in the weighted liberalization index for the values of the baseline migration rate available. One challenge to derive the standard errors and the confidence intervals displayed in Figure 1 is that $\text{Cov}(\hat{\beta}_1, \hat{\beta}_2)$ is not reported in GT22. However, as both $\hat{\sigma}_{\beta_1}$ and $\hat{\sigma}_{\beta_2}$ are reported in GT22, we overcome this challenge by replacing $\text{Cov}(\hat{\beta}_1, \hat{\beta}_2)$ in Eq. (4) with $\hat{\sigma}_{\beta_1}\hat{\sigma}_{\beta_2}\text{corr}(\hat{\beta}_1, \hat{\beta}_2)$, and report confidence intervals corresponding to the extreme values of $\text{corr}(\hat{\beta}_1, \hat{\beta}_2)$, i.e., -1 (darker shading) and 1 (lighter shading). Because $\text{corr}(\hat{\beta}_1, \hat{\beta}_2) \in [-1, 1]$, the actual confidence intervals always lie between these two extremes.¹¹

For the first specification in Panel A, the differential effect of a one standard deviation

⁹We use information on the baseline migration rate at the province level because information at the municipality level is much more limited in GT22; this drives a wedge with the econometric analysis, which is at the municipal-year level, but this is immaterial as our objective is just to compute the effect of a one standard deviation increase in the weighted liberalization index for various values of the baseline migration rate.

¹⁰More precisely, $\hat{\beta}_1 = 0.002$ (s.e. 0.001) and $\hat{\beta}_2 = -1.015$ (s.e. 0.331) in Panel A, and $\hat{\beta}_1 = 0.003$ (s.e. 0.005) and $\hat{\beta}_2 = -1.819$ (s.e. 0.572) in Panel B.

¹¹Note that $\hat{\sigma}_{\Delta Y_p}$ is bounded between $|\hat{\sigma}_{\beta_1} - \text{MigRate}_{p,t=0} \times \hat{\sigma}_{\beta_2}|$ (when $\text{corr}(\hat{\beta}_1, \hat{\beta}_2) = -1$) and $\hat{\sigma}_{\beta_1} + \text{MigRate}_{p,t=0} \times \hat{\sigma}_{\beta_2}$ (when $\text{corr}(\hat{\beta}_1, \hat{\beta}_2) = 1$), and that $\hat{\sigma}_{\Delta Y_p}$ monotonically increases with $\text{corr}(\hat{\beta}_1, \hat{\beta}_2)$. The lower bound equals zero when $\text{MigRate}_{p,t=0} = \hat{\sigma}_{\beta_1}/\hat{\sigma}_{\beta_2}$, i.e., $0.001/0.331 \approx 0.30$ percent for the specification in Panel A. This, in turn, explains why the smaller 95 percent confidence intervals for the specification in Panel A are hardly visible in Figure 1 for values of the migration rate corresponding to the seventy-fifth (0.31 percent) and to the eightieth percentile (0.38 percent).

increase in the weighted liberalization index stands at $-1.015 \times 0.0025 = -0.25$ percentage points (pp), and the corresponding t -statistic is equal to $-1.015/0.331 = -3.07$. In contrast, the average effect is close to zero (-0.02 pp) and far from significant (the t -statistic lies between -0.13 and -0.86 , depending on the value of $\text{corr}(\hat{\beta}_1, \hat{\beta}_2)$).¹² Thus, for this specification, the average effect is more than 10 times smaller in magnitude than the differential effect. Assuming $\text{corr}(\hat{\beta}_1, \hat{\beta}_2) = -1$, the province-specific effect is negative and significant for at most 40 percent of the provinces, i.e., 35 out of 84, representing 41.7 percent of the Filipino population in 1990.¹³ This share should be seen as conservative as it relies on coarse information on the distribution of the baseline migration rate and on the extreme assumption that $\text{corr}(\hat{\beta}_1, \hat{\beta}_2) = -1$. If we assume instead $\text{corr}(\hat{\beta}_1, \hat{\beta}_2) = 1$, the province-specific effect may be negative and significant for as few as one province.

Results are qualitatively similar for the second specification in Panel B. The differential effect is equal to -0.45 pp, and the corresponding t -statistic stands at -3.18 . The average effect is much smaller in absolute terms (-0.10 pp) and far from significance (the t -statistic is between -0.16 and -0.27). Assuming $\text{corr}(\hat{\beta}_1, \hat{\beta}_2) = -1$, the province-specific effect is negative and significant for at most 18 provinces (23.0 percent of the population). Assuming $\text{corr}(\hat{\beta}_1, \hat{\beta}_2) = 1$, it is negative and significant for no province.

4 Conclusion

We show that GT focus on a differential effect, and that this effect is insufficient to support the claim that more liberal reproductive health policies at destination reduce origin-community fertility. The actual values of $\hat{\beta}_1$, of its standard error, of $\text{Cov}(\hat{\beta}_1, \hat{\beta}_2)$, and of the baseline migration rate for the 84 Filipino provinces are necessary to derive the average and province-specific effects to support such claim. Our analysis using parsimonious information retrieved from an earlier working paper version highlights that the average effect is not significant, and that the province-specific effect is negative and significant in few provinces, even though the differential effect is negative and highly statistically significant.

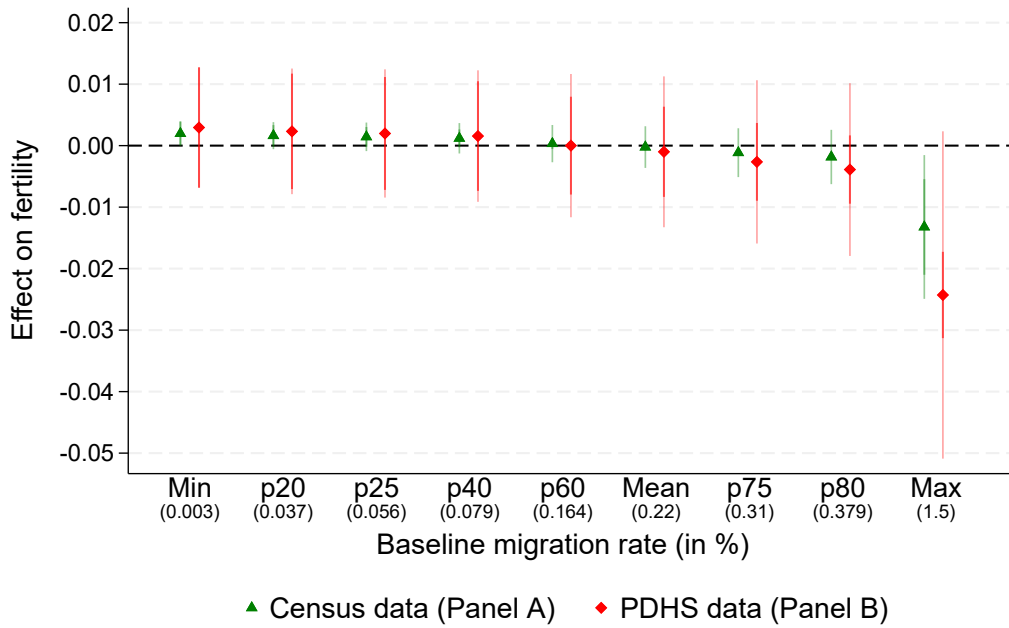
¹²The low values of the t -statistic imply that the conclusion about the lack of significance of the average effect is unaffected by the rounding of $\hat{\beta}_1$ and of its standard error $\hat{\sigma}_{\beta_1}$, for which only the last digit is different from zero.

¹³We draw on the 1990 census data from [Ruggles et al. \(2025\)](#) and on Figure 1 in GT22 to derive the share of the Filipino population living in the provinces belonging to each quintile of the distribution of the baseline migration rate.

References

- Finlay, Jocelyn E., David Canning, and June Y.T. Po**, “Reproductive Health Laws Around the World,” 2012. PGDA Working Papers 9612, Program on the Global Demography of Aging.
- Godlonton, Susan and Caroline Theoharides**, “Behavioral responses to temporary migration: An examination of origin-country fertility,” 2022. Available at: https://www.colorado.edu/economics/sites/default/files/attached-files/theoharides_0.pdf.
- **and** —, “Diffusion of Reproductive Health Behavior through International Migration: Effects on Origin-Country Fertility,” *American Economic Review*, 2025, 115 (10), 3597–3637.
- **and** —, “How international migration shapes fertility and reproductive health back home,” 2026. VoxDev Article published on January 28, 2026, available at: <https://voxdev.org/topic/migration-urbanisation/how-international-migration-shapes-fertility-and-reproductive-health>.
- Huntington-Klein, Nick, Claus C. Portner, Ian McCarthy, and The Many Economists Collaborative on Researcher Variation**, “The sources of researcher variation in economics,” 2025. NBER Working Paper 33729, National Bureau of Economic Research.
- Ruggles, Steven, Lara Cleveland, Rodrigo Lovaton, Sula Sarkar, Matthew Sobek, Derek Burk, Dan Ehrlich, Jane Lee, and Nate Merrill**, “Integrated Public Use Microdata Series, International: Version 7.6 [dataset],” 2025. Minneapolis, MN: IPUMS. <https://doi.org/10.18128/D020.V7.7>.

Figure 1: Effect of a one standard deviation increase in the weighted liberalization index on fertility at various points of the distribution of the baseline migration rate



Notes: This figure shows the effect of a one standard deviation increase in the weighted liberalization index on the fertility rate measured from the census (green triangle) and the PDHS (red diamond), at various points of the distribution of the province-specific baseline migration rate: $\hat{\beta}_1 + \hat{\beta}_2 \times MigRate_{p,t=0}$. Shaded bars show 95 percent confidence intervals. As we do not know $\text{corr}(\hat{\beta}_1, \hat{\beta}_2)$, we display two confidence intervals for each estimate: a conservative one assuming $\text{corr}(\hat{\beta}_1, \hat{\beta}_2) = 1$ (lighter shading); a less conservative one assuming $\text{corr}(\hat{\beta}_1, \hat{\beta}_2) = -1$ (darker shading). *Sources:* Authors' elaboration based on information reported in [Godlonton and Theoharides \(2022\)](#) and [Godlonton and Theoharides \(2025\)](#). Estimates of β_1 , β_2 , and their standard errors are from GT22 (Table 4). Baseline migration rate: min, mean, and max are from GT (Table A2); p25 and p75 are from GT (p. 3616); p20, p40, p60, and p80 are from GT22 (Figure 1).

Appendix A Relation to pre-registration

This comment builds on a note we pre-registered on OSF on February 11, 2026, when the replication package of the original article was still not available online. The analysis we pre-specified in that note relies on the distribution of the baseline migration rate, on $\hat{\beta}_1$, on its standard error, and on $\text{Cov}(\hat{\beta}_1, \hat{\beta}_2)$. As we were unable to recover these quantities from the replication package or from the Authors, we could not conduct our analyses as planned. Yet, a few deviations between this comment and the pre-registered note are worth highlighting:

1. Figure 1 shows the effect of a 1 SD increase in the weighted liberalization index on the fertility rate for several values of the baseline migration rate: min, p20, p25, p40, p60, mean, p75, p80, and max. Our pre-registered note specified that we would show this effect for each of the 84 provinces, however, we could not access information on their baseline migration rate. In addition, while we pre-registered that we would also report effects on contraceptive use, infant mortality, and maternal mortality, for these outcomes we lack estimates of β_1 to do so (even from the working paper).
2. In the pre-registered note, we did not pre-specify that we would use estimates from the working paper. We decided to use these estimates to provide evidence of the empirical relevance of our arguments and to relax otherwise binding data constraints.
3. In the pre-registered note, we did not pre-specify that we would examine the average value of ΔY_p . This is a natural quantity of interest that became apparent while working on this comment.
4. A large part of the pre-registered note is devoted to show that “Coefficient $\times$ *IQR*” corresponds neither to (i) the variation in the dependent variable that is induced by a simultaneous increase in both interacted variables, nor to (ii) the variation in the dependent variable that is induced by an increase in the weighted liberalization index alone. We then pre-specified the analysis we would conduct to examine these quantities. In this comment, we decided to focus on (ii), as (i) depends crucially on the direct effect of the baseline migration rate, which is absorbed by province fixed effects and therefore impossible to identify. In the note, we had introduced the (necessary) assumption that this unidentified coefficient was equal to 0.
5. Our pre-registered note also considered the relationship between the quantity “Coefficient $\times$ *IQR*” and the effect of a simultaneous increase in both interacted variables,

as the original article referred to “policy exposure” as the product of the baseline migration rate and of the weighted liberalization index. In preparing this comment, we decided to drop this possible interpretation of the quantity reported in the original article, as we noticed conflicting use of the expression “policy exposure”. As mentioned in footnote 3 of our comment, GT use the expression “policy exposure” to refer both to the weighted liberalization index (e.g., at pp. 3608 and 3610) and to its interaction with the baseline migration rate (e.g., at pp. 3609 and 3614).